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An Analysis of Unemployment Insurance Durations Since the 1990-1992 Recession

Final Report

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Karen E. Needels
Walter Nicholson

Submitted to:

U.S. Department of Labor, ETA/UIS
Rm. S-4231, 200 Constitution Ave.
Washington, DC 20210

Project Officer:
Brenda Bruun

Submitted by:

Mathematica Policy Research, Inc.
P.O. Box 2393
Princeton, NJ 08543-2393
(609) 799-3535

Project Director:
Karen Needels

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Karen Needels
Project Director

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EXECUTIVE SUMMARY

The average duration of insured unemployment has remained high since the end of the most recent recession, despite lower unemployment rates generally. Overall, the estimates presented in this report suggest that average durations increased by between 1.1 and 1.4 weeks in the post-1992 period relative to what might have been predicted based on historical data. This figure represents approximately a nine percent increase in the average duration for which unemployment insurance (UI) benefits are paid.

Increased average UI durations may be of concern to policymakers, for several reasons. To the extent that they represent increasing labor market difficulties that specific types of workers are facing, increasing average durations may suggest the need for new labor market initiatives to help those workers find new jobs. The increases may also reflect hardships that certain categories of unemployed workers are facing, even in the current "full-employment" economy. Policymakers may wish to consider ways in which UI policy (or, possibly, other income maintenance policy) might be adjusted to meet these needs. Finally, because increases in average UI durations imply increased aggregate levels of benefit payment under the program, these findings may raise concern about the adequacy of current UI trust fund levels. The present report, however, focuses primarily on identifying the reasons that average UI durations have increased relative to historical norms without explicitly addressing these larger policy concerns.

The review of the literature on UI durations presented in this report suggests several potential reasons for the recent increases, including (1) changes in UI laws that affect duration, (2) changes in the geographic distribution of claimants among the states, and (3) changes in the composition of the unemployed population that tend to favor longer durations. To assess the relative importance of these effects, the report contains a detailed analysis of aggregate data at both the national and state levels. It also includes an examination of claimant-level data, from four states, that seek to identify possible effects that may have been obscured in the aggregate statistics. The general conclusion of the analysis is that most of the increase in average UI durations is coming from the labor market itself (most notably from the increased average length of workers' unemployment spells), not from changes in UI policy. Specifically, the analysis presented here concludes that:

- ***Several factors related to the labor market appear to be the most likely explanations for the observed increase in average UI durations:***

  - Recent trends in the average duration of unemployment play an important role in explaining why average UI durations are higher than might have been expected. As measured by the total unemployment rate, labor markets appeared to be quite healthy in the post-1992 period. However, the lengths of unemployment spells were longer than have usually been associated with such low unemployment rates; these longer lengths explain a large portion of the increase in average UI duration compared to historical patterns.

- Increases in the fraction of claimants in demographic groups who are likely to experience long unemployment spells (older workers, females, African Americans) have played an important role in lengthening average UI durations. This trend is especially visible in the claimant-level data.

- Changes in the industrial composition of the labor force, most notably the decline in manufacturing jobs, also seem to have played an important role in increasing average UI durations. This effect probably arises because manufacturing unemployment itself is usually associated with higher recall probabilities and shorter associated spells of compensated unemployment than other types of layoffs.

- ***Several other factors do not appear to explain increases in average UI durations:***

  - The aggregate analysis concludes that changes in weekly benefit amounts or in average potential durations at the state level cannot explain the increase in average UI durations relative to historical patterns.

  - Changing rates of UI recipiency (as measured by the ratio of the insured to the total unemployment rate) do not explain increasing average UI durations. Indeed, the estimates reported here suggest that average UI durations should have *decreased* in response to recent declines in the average rate of UI recipiency.

  - Changes in the relative share of UI caseloads among the states do not explain recent increases in average UI durations relative to historical experience.

In addition (although examining other sources of income for claimants' households was not an explicit focus of this report), the literature review suggests that UI claimants do not easily increase other family income rapidly in response to unemployment. Only small percentages of claimants collect other government transfers during UI benefit receipt, and there is no evidence that spouses' employment rates or earnings increased after the claimants became unemployed. Therefore, UI benefits are a major source of short-term income support for workers who collect them.

I. INTRODUCTION AND BACKGROUND

The average duration of insured unemployment has remained higher in recent years than would be expected on the basis of historical data, despite low unemployment rates generally. For example, the 1997 national figure for average duration exceeded 14 weeks, about the level of the late 1980s, when unemployment rates were higher. The 1997 figure also exceeds by one to two weeks the figures recorded in the early 1970s, when unemployment rates were below five percent. Some portion of this higher average unemployment insurance (UI) duration may be explained by features unique to the recession of the early 1990s and the subsequent recovery, but other forces may be operating as well. Generally, an investigation of UI durations may shed light on the sources of this trend, such as changes in rates of permanent job loss, changes in UI laws, or other sources.

The implications of an increase (compared to historical standards) in the average UI duration for individuals, the UI system, and the economy differ, depending on the reason for the increase. For example, if the increase occurs because of changes in UI laws, it may be that unemployed workers face greater disincentives to reemployment; unemployed workers, however, may have greater ability to develop skills or search for jobs that use their skills more efficiently since more generous UI benefits may cushion the financial strain caused by unemployment. Policymakers would need to decide whether having more generous UI laws is an appropriate allocation of resources. If, on the other hand, the increase in average UI duration is because of a change in the distribution of claimants across states, then it may be that no policy change is necessary, since, within certain guidelines established at the federal level, states control the characteristics of their own UI programs.

Alternatively, the increase in average UI durations compared to historical expectations may be attributable to structural changes in the labor market. If the fraction of all unemployed workers who are permanently separated from their former employers increases, average unemployment duration would be expected to increase. Other changes in claimants' demographic or economic characteristics, such as changes in the industrial or occupational composition of claimants, could also affect UI durations. If labor market changes are the cause, policymakers may want to consider changing the type of services available to unemployed workers to respond to a greater need for retraining and increased difficulty finding jobs.

Regardless of the reason, if the average UI duration is increasing compared to what would be expected, this pattern has implications for the UI system. Total dollars paid in benefits is the number of first payments times the average weekly benefit amount (WBA) times the average UI duration. For a given number of first payments and an average WBA, an increase in the average UI duration will cause a short-term decline in UI trust fund reserves; therefore, UI tax rates must increase in response to the increase in benefits paid out. The UI system, however, may be able to play a role addressing any recent changes in claimants' needs, and thus in reducing UI durations, through the profiling system implemented.

A desire to quantify the magnitude and source of the change in UI durations, and to understand the policy implications, motivates this research project. This chapter of the report provides the research context for this study, specifically the literature relevant to understanding average UI durations. The second chapter contains an analysis of the annual pattern of average UI durations over time, both at the national level and across states. The third chapter contains an analysis of claimant-level data to assess the importance of changes in characteristics between the late 1980s and the mid-1990s. The final chapter discusses the policy implications of the findings.

A. LITERATURE RELATED TO UI DURATIONS

Much research has noted recent increases in average unemployment and UI durations compared to what would be expected based on historical experience (see, for example, Loungani and Trehan 1997; McMurrer and Chasanov 1995; and Baumol and Wolff 1998). This research has been based on individual-level data on unemployment spells or weeks of UI benefits collected. However, no known research examines the recent pattern of *aggregate* UI durations, constructed (using UI administrative data provided by the states to the U.S. Department of Labor) as the number of weeks of UI collected in a time period divided by number of first payments in that time period. Therefore, although the statistical properties of measures of duration constructed from individual-level data and from aggregate data may differ, research on the modeling, estimating, and interpreting of the determinants of *individuals'* durations of unemployment must be used as a guide for interpreting changes in aggregate UI durations. This individual-level research focus makes sense since the underpinnings of observed aggregate durations are individuals' behavioral responses to job loss. Only through understanding of individual behavior can appropriate policy responses--for the individual and for the economy--be designed. Therefore, in this literature review, the theoretical and empirical research on individual behavior and any factors that may affect the ability to draw inferences about individual behavior from aggregate measures of UI duration are discussed.

1. Factors Influencing UI Durations

The discussion of the literature related to UI durations begins with an overview of the theory of individual job search and unemployment duration and the empirical findings on the characteristics that influence unemployment durations. Next, the relationship between the UI system and unemployment is discussed.

a. Model of an Individual's Time to Reemployment

Standard models of job search assume that unemployed workers conduct their job search to maximize their lifetime expected well-being (usually called "utility"), which is a positive function of income and a negative function of time spent working (Burdett 1979; and Mortensen 1977).[1] In a simple version of the model, workers know the distribution of wages being offered by firms, but they do not know the wage offered by each company until they contact that company. Although job search models may be mathematically complex, they imply a simple rule for which job offer a worker should accept: at any point in time, accept the first offer of a wage higher than some minimum acceptable wage, called the "reservation wage."

The reservation wage for each worker in each time period is the wage for which the expected lifetime stream of utility from accepting a job is equal to the expected lifetime stream of utility from remaining unemployed. As a practical matter, the reservation wage is usually a function of an individual's economic and demographic characteristics, such as education level, work experience, and other family income. In more complex versions of the model, the effects of characteristics may vary over time, and the reservation wage may vary over the worker's unemployment spell.[2] For example, a worker may update the expectations about the distribution of wages available, or he or she may be more willing to take a lower wage as savings are depleted.

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[1] An alternative model of the determinants of unemployment duration explicitly considers the trade-offs between labor market and leisure by constructing a budget constraint so that utility is maximized (see, for example, Moffitt and Nicholson 1982; and Decker 1997). Changes in UI program parameters or other factors affecting expected income or the value of leisure change the shape of the budget constraint and may, in turn, change the duration of unemployment.

[2] Several other model extensions have been developed, such as by Mortensen (1977), who allows for the possibility of layoffs in subsequent jobs, and Rogers (1998), who allows for claimants to update their expectations of UI entitlements because of possible benefit extensions.

Several demographic and economic characteristics have consistently been associated with the duration of unemployment: these include recall status, unionization, industry, other income in household, sex, marital status, availability and characteristics of UI benefits, and economic conditions (see, for example, Corson and Dynarski 1990; and Corson et al. 1999).[3,4] Whether an individual expects to be recalled to his or her former employer is one of the most important characteristics predictive of the length of the unemployment spell. In adapting the standard job search model, Katz (1986) and others have allowed workers to have the possibility of recall from their previous employer. Using Panel Study of Income Dynamics (PSID) data, he finds that the likelihood of finding a new job decreases as the perceived probability of recall increases. Workers with a stronger job attachment are more likely to leave unemployment quickly (by returning to their former employers) than are unemployed workers who are not job-attached (and must search for new jobs to exit unemployment) (see, for example, Brewster et al. 1978; Corson et al. 1977; and Corson and Dynarski 1990).[5]

Other demographic and economic characteristics, as well as characteristics of the prior job, have been found to be associated with the length of the unemployment spell (see, for example, Corson et al. 1999). Having lower education is associated with a longer time to reemployment, because of the worker's having fewer work-related skills. Being a married female is associated with a longer time

[3]The effects of the UI program and the business cycle on unemployment durations are discussed in more detail in subsequent sections.

[4]Another way to examine the factors influencing the time to reemployment is to look at the factors influencing the rate at which individuals leave unemployment, often called the "hazard rate" or "exit rate," and how the exit rate changes as the time unemployed increases. However, this approach is simply a transformation of the approach that looks at time to reemployment.

[5]Katz and Meyer (1990) also point out that the rate at which workers find new jobs over time may appear to increase in the aggregate if factors that raise the recall rate lower the rate of finding new jobs. That is, as workers who are recalled leave unemployment, those remaining unemployed may be more likely to find new jobs than the group who were expecting recall.

to reemployment, probably because of a higher premium on nonemployment activities or larger sources of outside income. Being African American and older is also associated with a longer time to reemployment, while having a prior job in manufacturing and being unionized are typically associated with a shorter time to reemployment. Higher unemployment rates are also associated with lower exit rates from unemployment and longer times to reemployment for individuals, presumably because fewer jobs are available (see, for example, Dynarski and Sheffrin 1990; and Katz and Meyer 1990).

b. The Relationship Between the UI System and Unemployment

Theoretical models focus on unemployment duration and not UI duration per se. Standard theory predicts that UI program parameters, such as the WBA and potential benefits duration, affect unemployment. Hence, measures such as the replacement rate (the WBA divided by some measure of prior or average earnings) are included as explanatory variables when researchers try to identify the determinants of individuals' unemployment durations.

Effects of the Potential UI Duration on Unemployment Duration. One of the key research questions pertaining to the UI system is: "How many weeks does extra potential UI duration add to the time to reemployment?" This research is important for understanding not only the disincentive effects of the regular UI system, but also the disincentive effects of benefits made available through the permanent extended benefits (EB) programs or emergency benefits programs such as Federal Supplemental Compensation (FSC), Federal Supplemental Benefits (FSB), and, most recently, Emergency Unemployment Compensation (EUC). The potential delays in reemployment (or increases in the total number of weeks of UI benefits collected) associated with providing extra weeks of benefits must be counterbalanced with potential advantages, such as the provision of extra income as workers' families face reductions in their earnings and of extra time so that workers may conduct a more thorough job search or participate in skills-developing activities, such as training.[6,7] Nevertheless, numerous attempts have been made to quantify the disincentive effects associated with the provision of UI benefits.

In their recent reviews of the literature on UI program disincentives, both Woodbury and Rubin (1997) and Decker (1997) conclude that estimates of the disincentive effects of extra UI benefits vary widely. By considering the econometric methods and the data sources used, Woodbury and Rubin conclude that the most reliable estimate of how much an extra week of benefits increases the expected duration of unemployment is 0.2 week or less.[8] Woodbury and Rubin, as well as Decker, emphasize that different workers may respond differently to the availability of extra benefits. For example,

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[6]Mortensen (1977) and Katz and Meyer (1990) point out that increasing either the WBA or the duration of UI benefits has two effects: (1) increasing the value of remaining unemployed, and (2) increasing the value of taking a job from which one may be laid off in the future. These researchers postulate, and the empirical data support the view, that the first effect is larger than the second.

[7]The research that has examined the potential disincentive effects of higher WBAs is not discussed here. However, Katz and Meyer (1990) conclude that the disincentive effects of potential duration are greater than those for the WBA. An increase in the potential duration of benefits with a budget-neutral decrease in the WBA would increase the disincentive effects of the UI system, although it would provide greater protection for the long-term unemployed. Woodbury and Rubin (1997) argue that short spells of unemployment are overinsured and long spells are underinsured, so workers who are averse to risk would prefer reductions in the WBA and budget-neutral increases in potential duration.

[8]The sample over which estimates (of UI program disincentive effects) are calculated may be important. Levine (1993) has analyzed how UI program disincentive effects may affect the unemployment durations of workers who do not receive UI. By potentially reducing the work search efforts of UI recipients, nonrecipients may find jobs more quickly. Levine (1993) calculates that increases in UI program generosity (specifically, through increasing the wage replacement rate) may, on net, decrease the unemployment rate because of shorter unemployment spells by nonrecipients. In addition, the UI program may affect transitions into and out of the labor force. Several reemployment bonus demonstrations explored whether lump-sum benefits to workers who become reemployed quickly would reduce their time to reemployment and the cost to the UI trust funds (Spiegelman and Woodbury 1987; and Decker and O'Leary 1995). The bonuses weakened the approximately linear relationship between weeks of UI collected and total benefits collected. However, the results of the most recent demonstrations were discouraging, in that reemployment bonuses are unlikely to generate net savings to the UI system, so it is unlikely that this type of incentive scheme will be used in the future to reduce average UI durations.

workers who expect recall may not delay their time to reemployment when extra benefits are available, whereas workers who are permanently separated from an employer may delay finding a new job (Corson and Dynarski 1990).

Woodbury and Rubin also point out that most of the research uses UI spells (rather than unemployment spells) to estimate UI disincentives, although this measure is an imperfect substitute for the actual duration of unemployment, because individuals' length of time unemployed after receiving their last UI check may vary considerably. Once again, it is noted that UI duration is typically used in research as a proxy for unemployment duration, and there has been little focus either on examining how UI duration differs from unemployment duration or on analyzing the unique properties of UI duration.

Modeling an Individual's UI Duration. A simple theoretical model of the observed UI duration for an individual can define UI duration as the minimum of the duration of the unemployment spell and the duration of potential UI benefit receipt.[9] If a claimant becomes reemployed before exhausting his or her UI entitlement, then that claimant will receive benefits only for the weeks prior to reemployment. In contrast, if a claimant takes a much longer time to become reemployed, then the number of weeks of benefits that claimant receives is constrained by his or her maximum potential duration of benefits.[10]

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[9]The potential duration of benefits is defined as the entitlement divided by the WBA, conditional on being determined eligible. If an individual who would be entitled to UI benefits if he or she applied chooses not to file for them, then the potential duration is zero.

[10]By design, UI claimants who are unemployed for a long time collect more benefits than claimants who are unemployed a shorter period of time. As discussed earlier, the availability of benefits may encourage claimants to remain unemployed until the time they exhaust benefits. In an attempt to weaken the UI disincentive effects, several reemployment bonus demonstrations explored whether lump-sum benefits to workers who become reemployed quickly would reduce their time to reemployment and the cost to the UI trust funds (Spiegelman and Woodbury 1987; and Decker and O'Leary 1995). The bonuses weakened the approximately linear relationship between weeks of UI collected and total benefits collected. However, the results of the most recent demonstrations were discouraging, in that reemployment bonuses are unlikely to generate net savings to the UI systems, so it is unlikely that this type of incentive scheme will be used in the future to reduce average UI durations.

Duration of the unemployment spell is a function of the availability and potential duration of UI benefit receipt (because of the disincentive effects discussed earlier) and other claimant characteristics, such as recall status, occupation, and industry (and others discussed above). Potential UI duration in turn is a function of the claimant's base period earnings (the level of earnings--which presumably depend on the claimant's characteristics and, possibly, the distribution of earnings across quarters) and state-specific UI program parameters (Woodbury and Rubin 1997).

Additional complexities can be added to this simple theoretical model, which assumes that the claimant receives weekly benefits without interruption. For example, the model can be expanded to incorporate the effects of disqualifications for failure to meet the work search requirements or to participate in other mandatory activities, the effects of a waiting week, the possibility that claimants may not receive their full WBA each week because of earnings, temporary withdrawal from the labor force, or the possibility of more than one unemployment spell during a benefit year.[11]

Researchers have made some attempts to address these complexities. Swaim and Podgursky (1992), Portugal and Addison (1990), and Addison and Portugal (1987 and 1992) note that UI recipiency is a function of expected unemployment duration, since workers who expect to be unemployed for a very short period of time may not file for or receive benefits (particularly in states with a waiting week).[12] Thus, administrative delays or increases in waiting week requirements may reduce

[11]Some of the considerations suggest that workers have withdrawn from the labor force for a temporary period of time: for example, a worker who fails to satisfy the job search requirements for UI benefits might have done so. However, the relationship between satisfying UI program requirements and unemployment is imperfect. For example, states vary considerably in their requirements for continuing eligibility, and extended or EB programs have often set standards for program eligibility different from those for the regular UI system. Anderson (1997) provides a comprehensive review of issues surrounding continuing eligibility.

[12]In addition, Portugal and Addison (1990) find that the effect of the wage replacement rate on duration may also be overstated by failure to consider selection into the UI program by unemployed workers who are eligible for benefits.

the fraction of the unemployed who participate in the UI program. Conventional studies may overstate the average duration of unemployment by UI recipients while understating it for nonrecipients if the studies fail to take into account the waiting week and other administrative delays associated with UI benefits collection (Portugal and Addison 1990).

The UI benefits collection periods of some claimants stretch over a considerably longer period of time than their average potential duration. In these instances, the UI spell may be a poor proxy for the claimant's unemployment spell. To address this problem, some researchers, using UI administrative data, exclude claimants whose gaps between the dates of first and last payment are considerably larger than the benefits collected divided by the WBA. In doing so, these researchers attempt to exclude from the analysis claimants with interruptions in their unemployment spell (see, for example, Grossman 1989; and Corson et al. 1986). However, concern about the appropriateness of these types of sample exclusions is warranted. The reason for the discrepancy between the number of weeks between the first and last payment and the weeks of benefits collected at the full WBA often is not clear; for example, the interruption in benefits collection may occur because of new short-term employment, a temporary withdrawal from the labor force, or temporary UI program disqualifications. Another possibility is that UI benefits collection was not interrupted, but the claimant consistently collected less than the full WBA each week because of earnings that reduced weekly benefits payments.

Some researchers have used these types of restrictions on their sample in an attempt to improve the correlation between the observed UI spell and the unobserved unemployment spell. However, workers excluded because of one or more of these potential reasons may be concentrated in a few industries or a specific part of the wage distribution, so systematic differences between the excluded and included groups may exist. These exclusionary restrictions may therefore bias inferences drawn about the nature of UI (or unemployment) spells and the relationship between spell length and explanatory variables, since analysis is conducted only on the included group.

Overall, however, many factors may mitigate the close statistical relationship between the duration of an unemployment spell and the duration of the UI spell. Unfortunately, many of these factors are hard to observe empirically. Some of the necessary data, such as on UI disqualifications or weekly payment amounts, may require complicated extractions from UI administrative records. Others--such as details about unemployment status after an individual exhausted benefits--may in principle be available from surveys of claimants (or unemployed individuals generally). However, collecting the details necessary may pose considerable logistical challenges because of the need for information on weekly activities or because the respondents may have difficulty remembering specific details.[13] Hence, although many researchers treat unemployment and UI durations almost interchangeably because of data limitations and because they are closely related, there is no consensus about *how* close unemployment and UI durations relate to each other for individuals.

[13]For example, surveys of UI beneficiaries have difficulty identifying whether individuals have changed labor force status one or more times between benefits exhaustion and being interviewed (see, for example, Brewster et al. 1978; and Corson et al. 1977). Swaim and Podgursky (1992) and Addison and Portugal (1992) discuss the empirical difficulties associated with identifying the durations of unemployment spells (and whether these spells are right-censored) when one has data on labor force status at a point in time rather than data on labor market status since the unemployment spell began. Specifically, they identify trade-offs between including or excluding different categories of displaced workers from the analysis on the basis of their labor force status at the time of the Displaced Worker Survey, which occurred up to five years after the job separation of interest.

2. Changes in the Composition of UI Claimants

As discussed earlier, a large body of literature on the factors influencing individual unemployment durations exists. However, even if individuals have unemployment spells that are not sensitive to the business cycle, aggregate data may reflect a change in the composition of the unemployed over the business cycle. Changes in the characteristics of the population of the unemployed (or UI claimants) may limit the ability to draw inferences from aggregate data on individual behavior. Typically, analysis of changes in the composition of UI claimants has focused either on changes attributable to the business cycle or on secular changes that have occurred independent of the business cycle. Each of these is discussed in turn.

a. Business Cycle Changes

Although cyclical downturns are often defined by worsening labor market conditions, such as increased unemployment through job loss, it is not clear theoretically whether business cycle downturns will be associated with increased or decreased unemployment (or UI) durations. On the one hand, downturns may be associated with increased durations, since workers will have a more difficult time finding employment as companies shed workers and deplete inventories. Any given worker would be expected to have a harder time finding a new job during an economic downturn: the frequency of job offers is expected to be lower during downturns, and the distribution of wages offered to workers may be less favorable.[14] Workers may take some time to adjust their reservation wages, either because of

[14] Two effects of a faster job arrival rate on time to reemployment exist: workers have more chances to exit unemployment, but they may be more selective in the jobs they choose. Several researchers, such as Burdett and Ondrich (1985), examined the shapes of wage offer distributions that would allow the net effect of a faster job arrival rate to be a shorter time to reemployment. Van den Berg (1994) furthered this research and concluded that the range of wage offer distributions that allows for an increased job offer arrival rate to generate a faster exit from unemployment on net is quite broad.

imperfect information about how wages have changed or because of inflexibility in the wages workers will accept (Hall 1995). However, this phenomenon reflects not a change in the composition of workers per se, but a temporarily decreased demand for labor.

The average duration of unemployment (and UI) may change also because of a change in the composition of unemployed workers over the business cycle. Because the fraction of unemployed workers who are on temporary layoff typically increases during downturns, and since workers on temporary layoff tend to have shorter unemployment spells than workers who are permanently separated from their employers, the average duration may decrease (Lilien 1982).[15] As the economy recovers, temporary workers are rehired and the pool of unemployed workers consists of a larger fraction of permanently separated workers.

Both of these potential effects--associated with changes in demand for specific workers and changes in the composition of the unemployed--suggest that UI (or unemployment) durations may increase or decrease in response to economic downturns. The relationship between the business cycle and durations must be determined empirically.

Empirical Findings from the 1970s and 1980s. Using different data sets during the 1970s and early 1980s, researchers generally have found that higher unemployment rates are associated with longer unemployment durations, although this has not uniformly been the case.[16] For example, Flinn and Heckman (1982) used a subset of the National Longitudinal Survey of Young Men from 1969 to 1971 to find that a higher (monthly) national unemployment rate is associated with longer unemployment spells. Katz (1986) and Dynarski and Sheffrin (1990) used PSID data from the early

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[15]Firms may respond to a business downturn by increasing the length of their layoffs, so the net effect may be smaller than if the composition were to change during a business boom.

[16]Rogers (1998) finds that the coefficients for national- and state-level unemployment rates have different signs on the exit rate from unemployment. She does not explore this finding in detail.

1980s, and Solon (1985) used UI claimants in Georgia from 1978 to 1979, to reach similar conclusions.[17]

Baker (1992) explicitly examines whether the change in duration over the business cycle is attributable to an increase in the incidence of unemployment, a change in the composition of the unemployed, or an increase in the unemployment durations within each category of workers. Using a synthetic panel data set constructed from monthly data from the Current Population Survey (CPS) outgoing rotation groups (1979 to 1988), Baker finds that both the duration of individual unemployment spells and the incidence of unemployment increase during economic downturns. However, the increase in duration of unemployment spells accounted for about 60 percent of the increase in the unemployment rate.

Baker finds little evidence to support the hypothesis that changes in the composition of workers can explain aggregate changes in unemployment duration over the business cycle. The shares of the unemployed accounted for by most subgroups of workers did not change significantly over the business cycle, except for prime-age males and subgroups by reason of unemployment.[18] Thus, he finds that the change in the composition of the unemployed over the business cycle cannot explain a significant portion of the change in the aggregate average duration of unemployment over the 1980s.

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[17]Katz (1986) uses the average annual county unemployment rate, whereas Dynarski and Sheffrin (1990) use the monthly national rate.

[18]Baker considered subgroups by reason for unemployment, by region, by sex and race, by sex and age, by sex and education, and by industry.

The Recession in the Early 1990s. A large body of literature has focused on how the most recent recession--officially from June 1990 to March 1991--differed from earlier recessions in both its causes and its effects on workers.[19] The consensus is that, in contrast to earlier recessions, this one was mild but the subsequent recovery was extremely slow. For example, long-term unemployment peaked 15 months after the official end of the 1990s recession, whereas it peaked about 6 months after the ends of the recessions of the 1970s and 1980s (Ilg 1994). The time between the peaks in many other labor market indicators (such as the number of involuntary part-time workers, the number of discouraged workers, and the number of permanent job losers) and the official end of the most recent recession was longer than the time between the peaks and the official ends of earlier recessions (Gardner 1994).

The composition of job losers also differed during the most recent recession: in terms of both their industries and occupations and their likelihood of returning to their former employers. Coming in part from industries and occupations (such as the professional and managerial occupations) that have historically escaped the effects of economic downturns, the unemployed during the recession of the early 1990s came from a much broader spectrum of the labor force than the unemployed during the recessions of the mid-1970s and the early 1980s. They also were much less likely to expect to return to their former employers: 86 percent of all job losers were permanent job losers, compared to 56 percent in earlier recessions (Gardner 1994).

Although the economy has been relatively strong for the past several years, the unusual shape of the recession in the early 1990s may explain some of the current pattern of average unemployment (or UI) duration. Hall (1995) models how the effects of job displacement at the beginning of a recession

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[19] See, for example, Blanchard (1993), Hall (1993), and Hansen and Prescott (1993) on the causes and Boisjoly and Duncan (1994), Gardner (1994), and Ilg (1994) on the effects on workers.

will linger. He finds that these initial job losses can explain new job losses that begin two years later, because experienced workers face greater probabilities of job loss resulting from their lower tenure levels at their new workplaces.

Although Hall does not examine this specifically within the context of current and previous recessions, an implication of his research is that a higher rate of permanent job loss by senior workers will have a ripple effect on subsequent unemployment.[20] Following the recession of the early 1990s, during which a greater fraction of the unemployed were permanently separated from their former employers, job loss and average unemployment duration would be at higher rates than what would typically be expected during the economic recovery. Nevertheless, the labor market effects of the past recession may be symptomatic of widespread changes in the labor market that have been occurring in the past 20 years. These are discussed in the next section.

b. Secular Changes in Composition of UI Claimants

As with the business cycle, secular changes in the composition of claimants--which are independent of the business cycle--may lead to changes in average UI duration. Two important recent secular changes are discussed: (1) the decline in the fraction of the unemployed who receive UI benefits, and (2) the change in the nature of employment relationships and job attachments.[21]

The UI recipiency rate (the fraction of the unemployed who receive UI benefits) began a gradual decline several decades ago, then dropped dramatically in the early 1980s, and remains low (McMurrer and Chasanov 1995; and U.S. Department of Labor 1998). Depending on the measure

[20]Hall (1995) also estimates the financial consequence of a job displacement, which results from fewer hours worked and lower earnings levels. He finds that the financial loss to a worker is about 120 percent of the worker's annual earnings, although this estimate will vary depending on how job losses attributable to displacement are defined.

[21]Although these two phenomena may be causally related, research has not typically integrated them.

used, the long-term drop was between 40 and 60 percent of the rate in the late 1940s, so that recipiency rates are now about 30 to 40 percent.[22,23] This decline hampers the UI system's ability to provide a temporary source of income support to unemployed workers and to act as an automatic stabilizer.[24]

At the same time, potential UI duration and average unemployment duration have generally increased (McMurrer and Chasanov 1995; and Woodbury and Rubin 1997). No known research has analyzed how the decline in UI recipiency has affected average UI durations. However, the average duration would change in response to compositional changes in recipients if different recipient groups have different average durations. For example, a shift of the unemployed from an area of high recipiency rates and high average duration, such as the Northeast, to an area with lower recipiency rates and lower average duration, such as the South or Southwest, would cause the average UI duration to decrease at the same time that the overall recipiency rate declines. Alternatively, a decline in recipiency could be associated with an increase in average durations. For example, if some workers are less likely than others to collect UI but more likely to collect for a longer period of time when they do collect, then recipiency rates could decline while average duration increases if such workers make up a greater fraction of the unemployed over time.

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[22] The recipiency rate tends to increase during recessions as the fraction of job losers among the unemployed increases.

[23] Vroman (1998) points out that the ratio of insured unemployment to total unemployment increased slightly since 1986, although it still remains low compared to historical rates.

[24] Researchers have used slightly different measures for UI recipiency, such as the ratio of the insured unemployment rate (IUR) to the total unemployment rate (TUR) or the ratio of UI claimants to total number of unemployed workers. As McMurrer and Chasanov (1995) point out, measures differ slightly in the timing of how they are measured, but they are highly correlated. Analysis of their patterns over time leads to the same conclusions.

The Decline in the UI Recipiency Rate. Several studies have tried to decompose the decline in UI recipiency (particularly during the 1980s) into different sources.[25] Changes in the labor market that have been considered as potential sources for the change in UI recipiency have been shifts from manufacturing to service sector employment, the decline in the unionization rate, increased rates of female employment, increasing quasi-fixed costs of hiring new workers, and the increased use of part-time and contingent workers. In addition to changes in the labor force, changes in the aggregate characteristics of UI program participants may have arisen because of expanded coverage of the UI system and because of changes in federal requirements and state eligibility rules (such as the federal taxation of benefits and the decline in the real value of benefits), most notably during the 1980s, that reduce the percentage of the unemployed who receive benefits.

Blank and Card (1991), for example, conducted a detailed analysis of why the UI recipiency rate declined during the early 1980s. Using Current Population Survey (CPS) data, they find that the decline in the recipiency rate cannot be attributed either to changes in the eligibility-determining characteristics of unemployed workers or to changes in state UI laws governing eligibility.[26] Instead, they find that declining take-up rates among eligible unemployed workers is responsible for almost all of the decline. Blank and Card estimate that about half the decline in take-up rates results from shifts in the distribution of the unemployed across states (from the Northeast, which historically has had high take-up rates, to the South, which has had lower take-up rates), while the rest is attributable to changes

[25]Bassi and McMurrer (1997), Vroman (1991), and McMurrer and Chasanov (1995) also provide a thorough review of some of the studies discussed here.

[26]They estimate that tighter state eligibility rules reduced benefit receipt slightly, but that changes in the composition of workers increased eligibility slightly. On net, these effects could not explain the sharp decline in recipiency rates observed in the early 1980s.

in take-up rates within states. Lower unionization rates are the predominant cause of lower within-state take-up rates.[27]

Research by Vroman (1991) further investigates the decline in UI recipiency rates. Like Blank and Card, he points out that the recipiency rate experienced a sharp decline in the early 1980s. However, Vroman notes that the main factors that Blank and Card attribute this to did not experience sharp changes during the same time period. The shift of the population toward the southern states, the decline in the manufacturing industry's share of employment, and the decline in unionization rates have been occurring gradually over time and were not isolated to the early 1980s.[28]

Vroman was unable to explore the effects of changes in state UI program eligibility rules, but he points out that the timing of many of these changes coincides with the timing of the reduction in recipiency rates. Using data from supplements to the CPS in 1989 and 1990, he found that the most common reason that unemployed job losers did not apply for benefits was that they thought they were not eligible. Changes in UI program rules in the early 1980s and the gradual shift of the unemployed to areas where less is known about the UI program may have affected this rate. Consistent with this supposition, Corson and Nicholson (1988) found that changes in UI program eligibility rules were responsible for 40 percent of the decline in recipiency rates from 1980 to 1982.[29]

The disparate research findings suggest that several factors may be at work: the shift toward states with low UI take-up rates among the eligible population, the declines in unionization and manufacturing,

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[27]Other characteristics associated with differences in state take-up rates are higher wage replacement rates, UI disqualification rates, and UI coverage rates. However, changes in these characteristics do not explain the decline in take-up rates, since the relative patterns across states remained roughly constant over time.

[28]Vroman (1998) concludes that shifts in the labor force to geographic areas with low recipiency rates can explain part of the gradual decline in recipiency, but further research needs to be done to understand why those areas have historically had low recipiency rates.

[29]Blank and Card (1991) attribute very little of the decline to these changes.

and changes in state eligibility rules in the early 1980s. However, it is difficult to quantify the share that each factor is responsible for.

In conclusion, although consensus does not exist among researchers, several labor market and UI program factors can explain part of the decline in UI recipiency rates over the past several decades. The labor market trends that may be responsible for some of this decline do not appear to be reversing. Changes in UI program requirements may be used to reverse some portion of the decline in recipiency rates, but researchers are not clear why recipiency rates vary dramatically across states. Understanding geographic differences in recipiency rates is probably the most important next step to addressing the decline in recipiency. Identifying the sources of changes in the pool of unemployed workers who receive UI benefits can help explain the patterns in average benefit durations as well.

The Nature of Jobs and Job Separations. Particularly during the 1990s, common public perceptions are that permanent job loss and the percentage of all jobs that are of "poor quality" have increased, while the prevalence of "lifetime jobs" has decreased. Some researchers focus on changes in the nature of *job separations*, such as increased rates of permanent dislocations or direct measures of time unemployed or of UI benefits collection (see, for example, Baumol and Wolff 1998; Butler and McDonald 1986; Farber 1998; Kletzer 1998; Loungani and Trehan 1997; and Valletta 1996). Other researchers focus on changes in *jobs*, such as increases in nonstandard employment arrangements and declines in traditional ones (such as "lifetime" jobs for workers with more than five years of tenure) (see, for example, Hall 1982; Levenson 1996; and Vroman 1998).

The changing patterns of job loss and duration without work may be caused by technological change, increased variation in sectoral shocks, changes in domestic demand, increased international competition, or other reasons. Although researchers have not yet reached a consensus about the sources of increased permanent dislocation, a consensus is emerging that structural change in the labor market is affecting the nature of both employment (by making it less secure) and unemployment (by lengthening its duration). A greater number of permanent dislocations would be expected to affect measures of long-duration unemployment because of the lack of recalls and workers' potential need to re-train.

Henry Farber has conducted a series of studies (1997a, 1997b, and 1998) that look at long-term employment and job loss during the 1980s and the first half of the 1990s. He concludes that the fraction of workers with long durations on their jobs fell substantially after 1993, although the distribution of job durations remained relatively stable prior to then.[30] Men, particularly the less educated, were much less likely to report having a long-term job, whereas women were slightly more likely. In addition, despite the sustained economic expansion, he finds that the overall rate of job loss increased during the 1990s; about 15 percent of workers were displaced during 1993, 1994, and 1995. This increase was larger among highly educated workers, particularly because the position or shift was abolished, although job loss was still more prevalent for lower-educated workers (Farber 1997a; and Kletzer 1998). One mystery that arises in Farber's research is that dislocated workers are considerably more likely to report during the 1990s that their dislocation was for "other" reasons besides plant closings, slack work, or abolished positions or shifts. Farber could not explore the source for the increase in the prevalence of this "other" reason for dislocation. Nevertheless, the trend that a greater fraction of the unemployed are permanently separated from their former employers is clearly an important potential explanation for the observed increase in average UI durations.

[30] Valletta (1996) points out that average job duration is not the best measure of job security, since workers may be less likely to initiate job separations in a climate when employers are more likely to initiate permanent separations. Therefore, research should focus on whether the employer or employee initiates the separation, and whether that separation is permanent or not. Specifically, Valletta finds that temporary layoffs have become less common, and employers are increasingly likely to rely on permanent job separations.

In addition to potentially increased rates of permanent job separation, employers seem to be relying more on nonstandard employment relationships.[31] This pattern may likewise affect the UI system's ability to provide adequate levels or durations of benefits. Both Levenson (1996) and Vroman (1998) find that temporary employment has increased over the early 1990s (and earlier), while part-time employment has not. Vroman analyzes the changes in use of several types of nonstandard employment relationships (both temporary and part-time work, as well as other types), how growth in these types of relationships affects UI recipiency rates, and how policymakers may respond. By analyzing growth patterns, unemployment rates, and UI recipiency rates for categories of workers in these nonstandard arrangements, he concludes that the growth in these nonstandard work arrangements cannot explain much of the decline in UI recipiency rates in the past several decades.[32] Nevertheless, he points out ways that UI program provisions may prevent some of these workers from collecting benefits (even though they may meet monetary eligibility standards). For example, he notes that the requirement that beneficiaries be available and willing to accept full-time work limits the eligibility of many part-time workers, who are interested only in obtaining new part-time work. In addition, the disqualification of job leavers for the duration of their unemployment spell prohibits the collection of benefits by many workers with nonstandard work arrangements. Changes in these provisions may increase UI recipiency rates modestly.[33]

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[31] A common potential explanation for this shift is that firms are trying to avoid the quasi-fixed costs--such as health care benefits--associated with hiring new full-time permanent workers.

[32] The growth in part-time employment in the 1950s may explain part of the decline during that time period.

[33] Baumol and Wolff (1998) also discuss ways in which the UI system may respond to the increase in unemployment durations as a result of technological change. They recommend that the UI system be expanded to 39 weeks instead of the "regular" 26 weeks, that UI replacement rates be increased, and that the government concentrate efforts on retraining workers to address their lack of technological skills. However, Baumol and Wolff do not conduct an analysis of the budgetary implications of these suggestions.

(continued...)

In conclusion, a consensus exists that a greater fraction of the unemployed are permanently separated from their former employers, although little consensus exists about the reasons for the shift (Kletzer 1998). As discussed earlier, permanent job loss has been found to be strongly associated with increased UI durations, but little research has focused on how the increased prevalence of permanent job loss over time may affect aggregate levels of UI participation or UI durations and how policymakers should respond. In addition, although the increase in some types of nonstandard employment arrangements may limit the ability of some workers to collect UI benefits, research has not focused on the impact of growth in nonstandard arrangements on UI durations.

3. How the Long-term Unemployed Support Themselves

Because UI benefits are provided as time-limited resources to individuals and their families to tide them over while they look for work, benefit levels are set in most states to replace approximately half the earnings of claimants while they were employed. This level is designed to balance the need to provide a financial cushion to workers so they can find a job that matches their skills with the need to ensure that workers look for employment while collecting UI benefits. If nothing else in the household changes, a claimant's household income will be lower while collecting UI than while the claimant was working. However, a claimant may be eligible for other types of benefits (such as other government transfer programs), and other household members may work and earn more in response to the claimant's unemployment.

A review of findings from prior research suggests that rates of receipt for means-tested cash benefits (such as welfare), means-tested in-kind benefits (such as food stamps), retirement benefits (such as social security and private pensions), and other benefits (such as workers' compensation) are

[33](...continued)

quite low, both before and during UI receipt (Smith and Vavrichek 1990; Corson and Dynarski 1990; Corson and Nicholson 1982; and Corson et al. 1999). In general, rates of receipt during unemployment increased slightly, but it is clear that these sources of income are insufficient to replace the income lost through unemployment.

Several studies found little evidence that UI claimants, and specifically exhaustees, were able to increase other family income rapidly in response to unemployment. Earnings from spouses or unmarried partners were an important source of earnings for recipients with a working spouse or unmarried partner, and poverty status was highly correlated with the absence of a spouse's income (Corson et al. 1999; and Smith and Vavrichek 1990).[34,35] There is no evidence that their employment rates or earnings increased after unemployment (Corson et al. 1999).

In summary, a review of other studies suggests that UI keeps a substantial portion of families from experiencing poverty-level incomes during the period of benefit collection. Other transfer payments and retirement benefits are not sufficient to keep families above the poverty level. The earnings of the spouse or unmarried partner were an important and sizable source of family income, but this source was available to less than half of recipients. The studies found no evidence of increased employment rates or earnings of the spouse/partner during the unemployment spell.

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[34]About three-fifths of claimants in the EUC study reported being married or living together unmarried, and about two-fifths reported that they had a spouse or unmarried partner who worked.

[35]The percentage of claimants during the EUC period whose household incomes were below the poverty line increased from about 12 percent before unemployment to 45 percent during benefit collection (Corson et al. 1999).

B. CONCLUSIONS

This literature review provides a context for the present study of why UI durations since the 1990-1992 recession are longer than is typical when unemployment rates are at the recently observed low levels. In summary, the reviewed research suggests there are four major determinants of UI durations: (1) the business cycle, (2) UI program laws, (3) claimant characteristics, and (4) characteristics of jobs and job separation. Since the current study is concerned with whether average UI duration is longer than would be expected after controlling for the business cycle, there are three remaining determinants of interest to this study. First, the literature review suggests how the characteristics of the UI program play a strong role in the duration on UI. Cross-sectional differences in program characteristics, such as average potential duration or the average WBA, explain some of the cross-sectional differences in duration. Second, the characteristics of claimants may differ from those of claimants in the past such that duration on UI is longer. For example, average duration may increase if claimants are more likely to be lower-educated or female than in the past. Third, the nature of jobs and job separation may have changed such that duration is longer. If workers who are entering unemployment--and the UI system--have lost their jobs for different reasons than in the past, then UI duration may increase. For example, if plant closings, company mergers, and downsizing have become increasingly common, workers may spend more time before becoming reemployed.

In all the studies reviewed, even those going back to the 1970s, recall status has been a key factor affecting the duration of UI (or unemployment, more generally). The literature finds that rates of job attachment during the 1990-1992 recession were lower than during prior recessions; this pattern may be continuing during the economic upturn. Research also suggests that increased use of alternative work arrangements may be changing the face of the labor market. Claimant characteristics and UI program characteristics may also be responsible for increased UI durations in the 1990s. However, the review of the literature suggests that these are less likely candidates for explaining the observed pattern, since some of the largest UI program changes occurred well before the recession of the early 1990s, and relatively few changes have occurred since then. Recent changes in claimant characteristics also have been relatively small, suggesting that these may explain only a small portion of the increase in average UI durations in the 1990s.

A review of the literature also suggests that UI benefits are an important component in keeping claimants' households above the poverty line, but they are often not large enough to do so without earnings from other household members. Income from government transfers, retirement benefits, and other household members' earnings do not significantly increase in response to the unemployment spell.

II. TIME SERIES ANALYSIS OF AVERAGE UI DURATION

The average duration of UI benefits has remained unusually high long after the end of the recession of the early 1990s, compared to what would be expected given historical patterns.[1] In this chapter, aggregate time series data (primarily from UI administrative records and from the Bureau of labor statistics) are used to quantify this longer duration and to examine possible reasons for the trend. The examination begins with a look at the data on the national level. It then shifts to a state-level analysis because of the additional details about labor market conditions and about possible changes in state UI laws that these disaggregated data provide. The general finding is that both the national and the state data support the conclusion that average duration of UI benefits during the period 1993-1996 was about 1.1 to 1.4 weeks higher than what might have been expected given the overall level of unemployment that prevailed and historical experience. Such an increase represented about a nine percent increase in average duration on a national basis. It appears that this increase can be attributed neither to changes in the distribution of UI claimants among the states nor to changes in the provisions of state UI laws. More likely, the results suggest that this increase was caused primarily by changes in the nature of the unemployment being experienced by UI recipients--especially the fact that the typical unemployed person experienced a longer duration of unemployment than had been true at similar stages of the business cycle in the past. It is also shown that other major labor

[1] Throughout this chapter, published figures on the average duration of unemployment benefits are used. These figures are calculated by dividing weeks of UI benefits paid during a period by UI first payments during that period. Thus the figures are not the averages of individual claimants' experiences. For annual data, this method of calculation may be fairly representative of what microdata would show. With quarterly or monthly data, however, problems raised by differences in the timing of the weeks paid and first payments series may be more severe. Such timing issues are discussed later in this chapter.

market trends, most notably the general decline in manufacturing employment and the related decline in the prevalence of short-term layoffs, may help to explain increased average UI durations.

A. NATIONAL ANALYSIS USING ANNUAL DATA

Figure II.1 reports national data on average UI durations over the period 1978-1996, which clearly shows that average durations seem to have been higher in the mid-1990s than during other periods of economic recovery.[2,3] Whereas these durations had consistently been around 13 weeks during years in which the economy operated at high levels of activity, in the period 1993-1996 average durations were at least 14 weeks in every year, and sometimes higher.

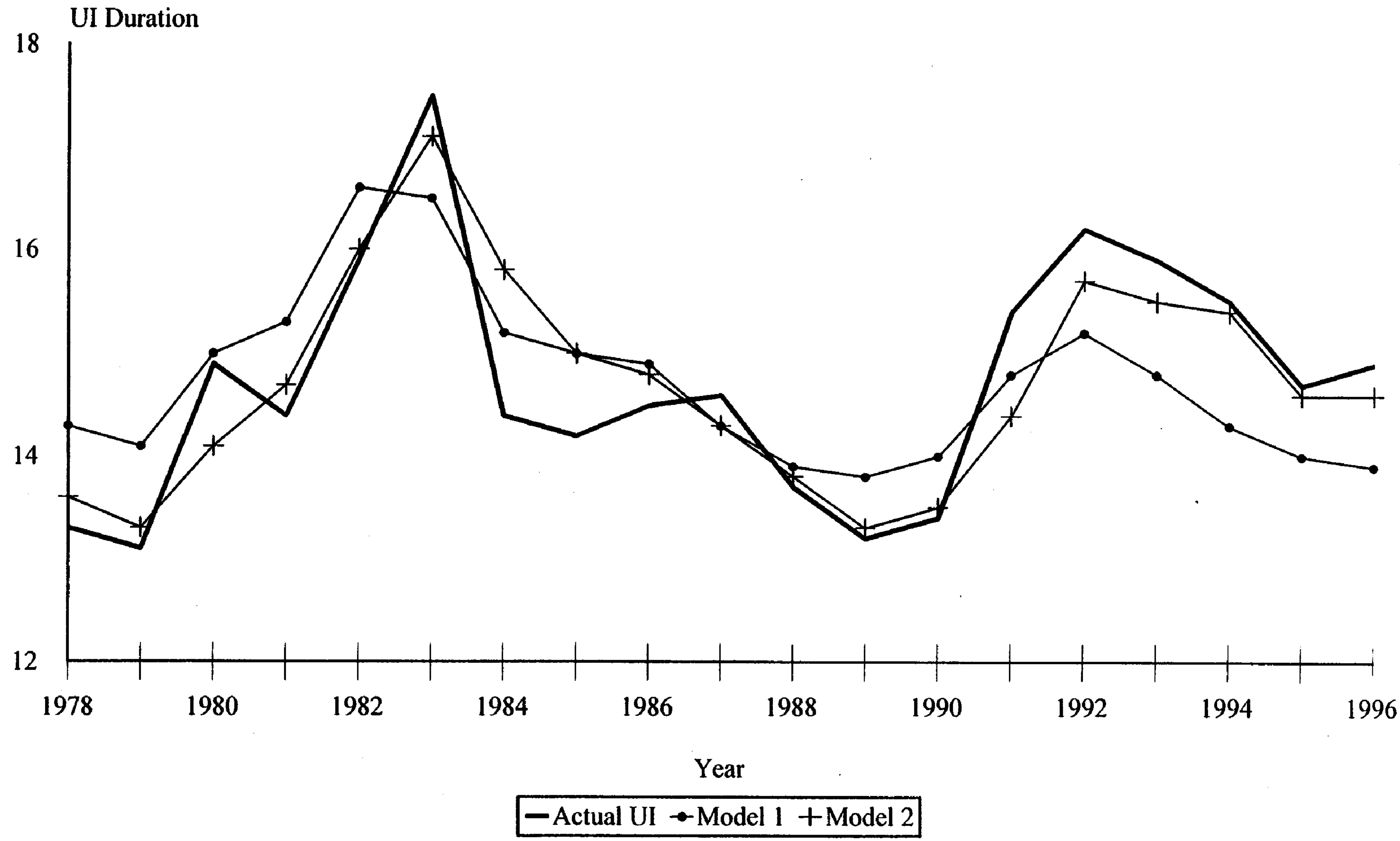
This apparent increase during the 1990s can be clarified further with simple regression analysis to control for the total unemployment rate (TUR) that prevailed at the time. Labeled "Model 1" in Figure II.1, these results forecast values of the average UI duration variable. (The specific regression equation underlying this forecast is reported in Table II.1, Equation 1.) These forecast values again clearly show the discrepancy of the mid 1990s: whereas forecasts based on the TUR track average UI duration fairly well through 1990, actual duration consistently exceeds forecast duration after that date. What is striking about the figure is the consistency of this discrepancy. Whereas prior to 1990 the forecast errors seem somewhat random (albeit with what appear to be substantial runs of positive and negative errors), the post-1990 forecast errors are consistently positive and all of about the same magnitude. This difference of a bit more than one week provides a visual hint that something may

[2]The data for Figure II.1 are shown in Appendix Table A.1.

[3]It is possible to use a much longer time period for these national data, but this shorter period was chosen so that the results would be directly comparable to the state results, which are constrained by the absence of unemployment rate data prior to 1978. An analysis of longer time periods using the national data suggests that the findings for the 1993-1996 period would be little changed if such a longer period were used.

FIGURE II.1

ACTUAL AND PREDICTED AVERAGE UI DURATIONS, 1978-1996



Note: Model 1 uses the TUR to predict UI duration. Model 2 uses the TUR and average unemployment duration.

TABLE II.1

NATIONAL REGRESSIONS, 1978-1996[a]

Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5	Equation 6
Total Unemployment Rate, Civilians 16 or Older	0.789*** (0.121)	0.548*** (0.159)	0.521** (0.258)	0.175 (0.346)	0.720*** (0.176)	0.562* (0.330)
Dummy Variable, =1 for Years 1993-1996	1.436*** (0.369)	0.671 (0.500)	0.904** (0.497)	0.663 (0.507)	1.200*** (0.360)	0.889 (0.560)
Duration of Unemployment, in Weeks		0.170** (0.081)		0.130 (0.091)		
Proportion of Total Employment in Construction			! 70.912 (84.026)	! 84.169 (81.343)		! 54.466 (100.897)
Proportion of Total Employment in Manufacturing			! 2.958 (9.658)	6.615 (11.442)		! 3.508 (14.223)
Proportion of Total Employment Who Are Females			! 13.365 (12.843)	! 20.731 (13.380)		! 9.977 (15.917)
Average Potential Duration for Regular Unemployment Benefits, in Weeks					! 1.779 (0.800)	! 0.802 (0.887)
Ratio of Average UI Benefit to Average Weekly Wage					10.650 (35.004)	3.261 (33.911)
Ratio of Insured Unemployment Rate to Total Unemployment Rate					2.075 (4.186)	1.957 (4.533)
Constant	9.078*** (0.858)	8.310*** (0.863)	21.612*** (6.811)	23.885*** (6.740)	47.656*** (18.957)	36.320* (20.586)

TABLE II.1 *(continued)*

Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5	Equation 6
R-squared	0.743	0.801	0.815	0.829	0.820	0.877
Standard Error	0.622	0.566	0.498	0.479	0.577	0.545
F	23.114***	20.087***	16.862***	15.532***	11.883***	8.900***
Durbin-Watson	1.675	1.690	2.418	2.392	2.368	2.588

[a]The dependent variable, average UI duration, has a mean of 14.72.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

have changed at some point just after the recession of the early 1990s (which officially ended early in 1991).

To examine whether this 1+ week difference was robust to alternative forecasting methods, a large number of national-level regressions on average UI duration were run that sought to control for factors that prior research had indicated might be important determinants of aggregate UI durations. In addition to the general measure of labor market strength (the TUR), these factors were grouped into two major categories: (1) variables that sought to characterize the unemployed population, and (2) variables that measure characteristics of the state and federal unemployment compensation system. In the first category, included measures consisted of the industrial composition of employment (because, for example, the decline in manufacturing employment may have reduced the incidence of short-term layoffs and increased average UI durations); the demographic composition of the unemployed (because increasing representation of women among the unemployed may have increased UI durations); and various measures of the duration of unemployment itself (because of the close connection that prior research has established between UI duration and unemployment duration). For the possible influence of UI policy, variables such as the average potential duration for which workers could collect UI benefits (which prior research has shown to increase *actual* UI durations), the UI wage-replacement ratio (which has also been shown to increase UI durations), and the ratio of the IUR to the TUR (a measure of the UI recipiency rate for which there were no strong prior beliefs about how the observed decline in this variable should have affected average UI durations) were used.

Table II.1 reports selected results.[4] The analysis focused particularly on whether the equations could explain the relatively high levels of average duration in the 1993-1996 period--that is, could other national-

[4]Variable definitions, means, and standard deviations are shown in Appendix Table A.2.

level variables reduce the estimated extent to which UI durations in the mid-1990s exceeded historical levels? Two conclusions are readily apparent in these results. First, the 1+ week discrepancy is robust to inclusion of variables representing characteristics of the UI system (such as the wage replacement rate or potential duration) or variables representing the industrial or demographic composition of the employment (such as the percentage of employees in manufacturing and construction).[5] Most of these variables also had estimated effects on average duration that were not statistically significant, in many cases because the variables themselves were highly correlated, making the estimation of independent influences very difficult with this sort of aggregated data.

The second major conclusion that can be drawn from the regression equations in Table II.1 is that one national-level variable--the measured average duration of *unemployment*--did have an important effect on the estimates. Whenever this variable was included in the regressions, the 1993-1996 discrepancy became statistically insignificant.[6] In Figure II.1, the line labeled "Model 2" uses both the TUR and average unemployment duration to forecast average UI duration. Here the forecasts track actual average UI duration fairly well even into the mid-1990s. Although this result is not surprising (for many of the unemployed, the duration of unemployment and the duration of UI benefits are probably identical), it does suggest strongly that the observed changes in average UI durations are probably being caused in large part by changes in the duration of unemployment itself. This fact, in turn, suggests that changes in the nature of unemployment itself (such as the changing relative importance of dislocated workers or of workers on

[5]The conclusion that the percent of workers in manufacturing has "no effect" on average UI duration did not hold up in the monthly analysis, which indicated that there may be some connection between the decline in manufacturing employment and increasing UI duration. See Sections C and D of this chapter.

[6]The estimated increase in average UI duration in the 1993-1996 period, although "not statistically different from zero" remained relatively large--approximately two-thirds of a week. See Table II.1, Equations 2 and 4.

short-term layoff) may be the ultimate cause of changing average UI duration. The high degree of correlation among the annual data series makes a precise estimation of such effects impossible, however.

B. NATIONAL ANALYSIS OF QUARTERLY DATA

A quarterly analysis allows us to gain further insights into the time series behavior of average UI duration by examining its seasonal variability and by exploring several macroeconomic indicators that are available on a quarterly basis. Table II.2 provides a summary of the findings, which closely mirror those given in Table II.1. Equation 1 shows that when the TUR is used as a cyclical indicator, average UI duration was approximately 1.3 weeks higher in the 1993-1996 period than would be expected given historical patterns.[7] The analysis explored whether each year in the 1993-1996 period needed to be separately estimated, but the hypothesis that all the effects in the years were identical could not be rejected.[8] Hence, the rest of the national estimates focused only on the average increase over the 1993-1996 period.

Equations 2 and 3 are representative of the estimates made with alternative measures of the business cycle (represented here by the rate of capacity utilization and the real gross domestic product [GDP] growth rate to explain changes in average UI durations). None of the other cyclical

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[7]Equation 1 (and all the other equations in Table II.2) included quarterly dummy variables that indicate that average duration is highly seasonal--average duration in the second quarter of the year is about 3.4 weeks higher than during other quarters. This finding probably represents differential seasonal timing in the two components of the average UI duration measure--weeks compensated during the second quarter, in part, represent first payments that are made early in the year, whereas first payments during the second quarter are usually at seasonally low levels. In the next section, it is shown that the monthly figures on average duration are affected even more significantly by such timing factors.

[8]The year-specific effects were 1.27 weeks (1993), 1.57 weeks (1994), 1.05 weeks (1995), and 1.49 weeks (1996).

TABLE II.2

NATIONAL QUARTERLY RESULTS
(DEPENDENT VARIABLE: AVERAGE UI DURATION)

Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5	Equation 6
Total Unemployment Rate	0.743*** (0.083)	--	--	0.745*** (0.088)	0.436*** (0.117)	-0.041 (0.172)
Rate of Capacity Utilization	--	-0.260*** (0.031)	--	--	--	--
Real Gross Domestic Product Growth	--	--	0.140*** (0.037)	--	--	--
1993-1996 Dummy Variable	1.346*** (0.295)	1.069*** (0.296)	0.802* (0.434)	1.391*** (0.450)	0.326 (0.403)	0.314 (0.334)
Ratio of Insured Unemployment Rate to Total Unemployment Rate	--	--	--	0.222 (1.144)	--	--
Average Duration of Unemployment (Weeks)	--	--	--	--	0.206*** (0.059)	--
Percent Unemployed Over 27 Weeks	--	--	--	--	--	2.386*** (0.474)
Constant	9.143*** (0.605)	35.381*** (2.500)	13.859*** (0.295)	9.014*** (0.891)	8.378*** (0.612)	12.147*** (0.805)
R-squared	0.749	0.738	0.604	0.751	0.777	0.801
Standard Error	1.058	1.080	1.352	1.099	1.001	0.947
F	58.42***	55.20***	28.42***	44.18***	56.39***	65.02***

NOTE: All equations also contained quarterly dummy variables. Data are from 1971 through 1996. The mean UI duration is 15.2 weeks. Standard errors are in parentheses below the coefficient estimates.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

indicators fit the data as well as did the TUR, as shown by the lower R-squared. However, some affected the estimated 1993-1996 increase in average UI duration--as is the case in Equation 3.[9] Further research (not reported) could detect no explicit pattern in the relationship between the cyclical measures used and the estimated mid-1990s increase in average UI durations. Therefore, it was concluded that the TUR is the best cyclical indicator for use in investigations of this type.

Equations 4 through 6 examined various labor market measures that might explain the observed increase in the average UI duration. As in Table II.1, the UI recipiency rate (as measured by the ratio of the IUR to the TUR) had a positive but insignificant effect on average UI duration, and the inclusion of this variable did not affect the coefficient of the 1993-1996 dummy. On the other hand, once any indicator of the duration of unemployment was included in the equations (average duration in Equation 5 and the percent unemployed over 27 weeks in Equation 6), the coefficient of the 1993-1996 dummy became statistically insignificant and one week smaller (0.3 week, compared to 1.3 weeks). This confirms the conclusion from the annual data that changes in average UI duration are being driven by changes in unemployment duration.

C. NATIONAL ANALYSIS OF MONTHLY DATA

Although many of the cyclical indicators used for the quarterly analysis are not available on a monthly basis, a variety of monthly analyses were conducted using the set of labor market variables that are available. Table II.3 presents a selection of the results. These are qualitatively similar to the

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[9]Inclusion of all three cyclical variables gave an estimated coefficient for the 1993-1996 dummy of 1.2 weeks.

TABLE II.3

NATIONAL MONTHLY RESULTS
(DEPENDENT VARIABLE: AVERAGE UI DURATION)

Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5	Equation 6
Total Unemployment Rate	0.737*** (0.055)	0.422*** (0.077)	0.286*** (0.078)	0.849*** (0.119)	0.742*** (0.057)	0.264*** (0.086)
1993-1996 Dummy Variable	1.341*** (0.195)	0.287 (0.264)	0.037 (0.249)	1.331*** (0.195)	1.289*** (0.298)	0.067 (0.325)
Average Duration of Unemployment (Weeks)	--	0.213*** (0.038)	--	--	--	0.338*** (0.048)
Percent Unemployed Over 27 Weeks	--	--	17.908*** (2.391)	--	--	--
Ratio of Insured Unemployment Rate to Total Unemployment Rate	—	—	—	—	0.058 (0.736)	3.764*** (0.859)
Percent of Unemployed on Layoff	--	--	--	-43.817 (41.357)	--	--
Constant	8.087*** (0.442)	7.288*** (0.445)	8.694*** (0.414)	7.773*** (0.532)	8.000*** (0.612)	4.744*** (0.730)
R-squared	0.817	0.835	0.846	0.818	0.816	0.845
Standard Error	1.209	1.152	1.110	1.208	1.233	1.136
F	102.65***	107.24***	116.95***	95.43***	85.41***	97.30***

NOTE: All equations also contained 11 monthly dummy variables. Data are from 1971 through 1996. The mean UI duration is 15.2 weeks. Standard errors are in parentheses below the coefficient estimates.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

annual and quarterly results described earlier.[10] Regressions that contain only the TUR as a cyclical indicator imply that UI durations averaged more than a week longer in the 1993-1996 period (Equations 1, 4, and 5). That differential is eliminated by any inclusion of variables measuring underlying unemployment durations (Equations 2, 3, and 6). However, inclusion of the layoff rate (which was expected to have a negative effect on average UI duration) did not adequately control for changes in unemployment durations.

Finally, Equation 6 illustrates a potentially important relationship between the average duration of unemployment and the rate of UI recipiency (here, the ratio of the IUR to the TUR). Inclusion of the IUR/TUR ratio increased the estimated effect of unemployment durations (compare Equations 2 and 6). The coefficient in Equation 6 is large enough so that the increase in average unemployment duration in the 1993-1996 period generated a bigger rise in average UI durations than was observed in the data.[11] However, some of this potential increase was mitigated by a decline in the recipiency rate. These recipiency rates have fallen during the 1993-1996 period, compared to historical levels, and therefore cannot explain the increase in average UI durations.

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[10]All of the equations reported in Table II.3 also controlled for the large month-to-month variation in the measured UI duration. The large monthly variability in the average UI duration figure arises from how it is calculated--as the ratio of two flows with major and differing monthly patterns. For example, measured UI durations are about three weeks below average in January (when UI first payments peak) and four weeks above average in the March-to-June period (when first payments are at low levels). Hence, considerable care must be taken to control for seasonality when seeking to discern trends in the monthly average UI duration data.

[11]Controlling for other factors, unemployment duration was about 4.5 weeks longer in the 1993-1996 period than would have been expected given the TUR. According to Equation 6, this change would have generated about a 1.5-week increase in the average UI duration. On the other hand, the observed 10 percent decline in UI recipiency (from 0.437 in the period 1978 to 1992 to 0.396 in the period 1993 to 1996) would have tended to reduce UI duration by about 0.4 week.

D. STATE ANALYSIS OF ANNUAL DATA

State-level data on UI durations was used to address three general questions that could not be addressed with the national data. First, these data were used to examine whether the national trends could be explained simply by changes in the composition of UI caseloads across the states. Next, it was asked whether the annual data that are available on some of the characteristics of state UI systems could help explain state-specific changes in UI duration. Finally, detailed monthly data at the state level were used to see whether the large increase in observations provided by such detailed data could add further insights into possible changes in the determinants of UI durations.[12]

As a first step in the analysis of state-level data on average UI durations, it was whether the national increase in average UI duration can be explained simply by the shifting composition of UI caseloads among the states. Because average UI duration tends to have substantial and lasting differences among the states, small changes in the relative sizes of states' UI programs could have had an important effect on the national data.[13] Table II.4 reports on national average UI duration using two different weighting schemes for aggregating across the states: (1) weights based on UI first payments during the 1978-1980 period, and (2) weights based on UI first payments during the 1994-1996 period.[14] The table illustrates these data for two comparison years with roughly similar labor market conditions: 1978 and 1996. Overall, the weighting

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[12]Unlike the national data, no data on the state level are available on a quarterly, but not a monthly, basis. Hence, a separate analysis of quarterly data at the state level was not undertaken.

[13]Alabama, Georgia, and Virginia had the lowest levels of average UI duration during the sample period (about 10 weeks), whereas other states (for example, Massachusetts, New Jersey, and Pennsylvania) had average levels above 16 weeks. The District of Columbia had an average UI duration of greater than 19 weeks during the sample period.

[14]Weighting schemes based on three-year averages in first payments were chosen to reduce the influences in year-to-year variations in states' first payments figures.

TABLE II.4

NATIONAL AVERAGE UI DURATION USING DIFFERENT STATE WEIGHTS

	Weights Using First Payments in	
Year	1978-1980	1994-1996
1978	13.20 weeks	13.17 weeks
1996	14.74 weeks	14.92 weeks

schemes made little difference. Weighted national average UI duration was about 13.2 weeks in 1978 and about 14.8 weeks in 1996, regardless of which weights were used. Use of other years and weighting schemes produced essentially the same results. Hence, it was concluded that the national increase in average UI duration was not being caused in any significant way by shifts in the composition of the UI caseload among the states.

A primary advantage of the state-level data is that they permit a more detailed investigation of whether changes in the provisions of state UI laws are responsible for changing UI durations. The principal shortcoming of these data, however, is that state-level information on the economic and demographic characteristics of the unemployed is not as rich as information available at the national level. Especially important, given the results from the national-level analysis, is the absence of state-level data on the duration of unemployment. This omission means that the results reported here should be viewed mainly as suggestive, since they have not controlled for an important determinant of average duration.

Table II.5 reports on a series of pooled regressions for 51 UI jurisdictions over the period 1978-1996.[15,16] These regressions used average UI duration as the dependent variable, and, because of the substantial cross-state differences in the levels of that variable, most of the equations were estimated using the "fixed-effects" estimation procedure.[17] By controlling for state-specific determinants of average UI duration, this procedure should provide better estimates of how variations in economic conditions and in UI policy variables across the states affect average UI duration.

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[15]That is, 50 states plus the District of Columbia.

[16]Variable definitions, means, and standard deviations are shown in Appendix Table A.5.

[17]This procedure amounts to including a time-invariant dummy variable for each state in the regressions. Similar results were obtained by using ordinary least squares and by using the "random- effects" method for estimating pooled regressions (Greene 1993).

TABLE II.5

POOLED REGRESSIONS, 51 UI JURISDICTIONS, 1978-1996[a]

Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5
Total Unemployment Rate, Civilians 16 or Older	0.691*** (0.022)	0.664*** (0.022)	0.671*** (0.022)	0.595*** (0.024)	0.691*** (0.022)
Dummy Variable, =1 for Years 1993-1996	1.147*** (0.089)	1.155*** (0.087)	1.294*** (0.088)	1.168*** (0.087)	1.146*** (0.089)
Average Potential Duration for Regular Unemployment Benefits, in Weeks		0.122*** (0.031)	0.107*** (0.030)	0.122*** (0.030)	
Ratio of Average UI Benefit to Average Weekly Wage		8.521*** (1.448)	5.975*** (1.462)	6.570*** (1.430)	
Ratio of Insured Unemployment Rate to Total Unemployment Rate			3.393*** (0.499)	3.604*** (0.488)	
Dummy Variable for Years in Which FSC or EUC Benefits Were Available				0.509*** (0.075)	
Constant					9.081*** (0.330)
R-squared	0.843	0.851	0.858	0.865	0.843
Standard Error	1.091	1.064	1.039	1.014	1.063
F	4914***	1737***	1379***	1166***	---

NOTE: Equations 1 through 4 were estimated using fixed effects; Equation 5 was estimated using random effects.

[a]The dependent variable, average UI duration, has a mean of 13.82.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

UI = unemployment insurance.
FSC = Federal Supplemental Compensation.
EUC = Emergency Unemployment Compensation.

The most basic regression (Equation 1) used only the state TUR as a single independent variable. This equation yielded results similar to the national estimates presented in the previous section. Overall, the equation suggests that, during the 1993-1996 period, the average state UI duration was about 1.1 weeks above what might have been expected based on the strength of state labor markets. The equation also implies that each 1 percentage point increase in the TUR resulted in an extra average UI duration of about 0.7 week--again a figure close to the national estimate (0.8).

Three aspects of state UI systems were found to have significant effects on average UI duration (Equations 2 through 4). The average potential duration (that is, the average length of regular unemployment benefits for which claimants are eligible) was estimated to have a significant positive impact on actual duration. Each one-week increase in average potential duration was estimated to increase the average UI duration by about 0.1 week--a figure close to estimates of the disincentive effects of additional potential duration found in many other studies (Woodbury and Rubin 1997). Similarly, generosity of a state's UI WBA (as measured by the ratio of the WBA to the average weekly wage in the state) was found to have a significant positive effect on average duration. Consistent with the estimates from other research (reviewed in Decker 1997), a 10 percentage point increase in this measure of generosity was estimated to increase average duration by between 0.6 and 0.9 week. Finally, the UI recipiency rate was found to have a small but significant positive effect on average duration--each 10 percentage point increase in this rate (again, as measured by the ratio of the IUR to the TUR) was associated with an increase in average UI duration of about 0.3 week.

Whether extended-benefits policy at the federal level had any effect on states' average UI duration was examined. This was done by looking at periods during which first payments were made under the two emergency programs: (1) the FSC program in the early 1980s, and (2) the EUC program in the

1990s.[18] As shown in Table II.5, Equation 4, results for this specification suggested that, other things being equal, availability of such emergency extensions increased average duration under regular UI programs by a statistically significant half week.

As Table II.5 shows, inclusion of all these UI-related variables had little effect on estimates of the increase in average UI duration in the 1993-1996 period. For the average potential duration and wage replacement variables, the explanation is simple: on average, neither variable exhibited any change (relative to its past values) during the 1993-1996 period. Hence, although these variables do affect average durations in general, they cannot explain the recent changes. Similarly, taking into account the availability of emergency extended benefits had little effect, because such benefits were paid only for a small portion of this period.

UI recipiency rates were estimated to be about four percentage points lower during the 1993-1996 period than might have been predicted on the basis of historical experience and the strength of local labor markets. According to these estimates, this fact actually *reduced* average UI durations by about 0.1 week from what they would have been had recipiency rates remained unchanged. Hence, as in the national estimates, this factor also cannot explain *increasing* average durations.

Table II.6 looks at the experiences of each state separately. The table reports the average UI duration for each state in 1996, together with two estimates of the extent to which the average duration in 1993-1996 exceeded historical levels. Each specification allowed each state to have a different coefficient for the 1993-1996 dummy variable. The first such estimate (referred to as "Common

[18]Periods when first payments were made under the regular EB program were also examined, but, because the data series on EB contained many missing values, these results are not reported here.

TABLE II.6

AVERAGE UI DURATION AND THE CHANGE IN AVERAGE DURATION, BY STATE

State	Average Duration in 1996	1993-1996 Change in Average Duration "Common Slope"	1993-1996 Change in Average Duration "Different Slopes"
Alabama	10.5	0.61	!0.39
Alaska	15.2	0.39	0.20
Arizona	14.5	1.23**	1.32**
Arkansas	12.1	1.79**	1.37**
California	16.9	0.81	0.60
Colorado	12.4	1.35**	1.47**
Connecticut	15.9	3.42**	3.11**
Delaware	16.9	2.43**	2.09**
District of Columbia	19.2	!0.48	!0.45
Florida	14.3	1.93**	1.92**
Georgia	9.6	0.55	0.47
Hawaii	17.7	2.64**	2.69**
Idaho	12	0.71	0.92
Illinois	17.1	0.99	0.93
Indiana	11.2	1.43**	1.22**
Iowa	12.5	0.38	0.01
Kansas	13.7	0.42	0.36
Kentucky	12.2	!0.94	!0.35
Louisiana	14.9	0.00	0.10
Maine	14.2	0.43	0.43
Maryland	15.7	1.89**	1.88**
Massachusetts	16.3	0.85	0.85
Michigan	11.3	0.72	0.41
Minnesota	14.3	0.76	0.64
Mississippi	13.8	2.17**	1.62**
Missouri	13.4	1.85**	1.52**
Montana	14	1.57**	1.08
Nebraska	11.8	0.46	0.57
Nevada	13.9	0.69	0.68
New Hampshire	9.8	2.22**	2.12**
New Jersey	17.4	1.15**	1.20**
New Mexico	16.4	0.83	0.74
New York	19.2	1.10**	1.10**

TABLE II.6 *(continued)*

State	Average Duration in 1996	1993-1996 Change in Average Duration "Common Slope"	1993-1996 Change in Average Duration "Different Slopes"
North Carolina	9.6	1.11**	1.08**
North Dakota	12.3	!0.67	!0.36
Ohio	13.6	1.42**	1.55**
Oklahoma	12.7	1.21**	1.33**
Oregon	15.3	2.53**	2.28**
Pennsylvania	16.8	1.79**	1.93**
Rhode Island	15.7	1.44**	1.49**
South Carolina	11.1	0.78	0.67
South Dakota	10.9	0.06	0.26
Tennessee	12.1	0.85	0.86
Texas	15.8	1.86**	1.92**
Utah	10.9	0.17	0.62
Vermont	14.4	1.47**	1.62**
Virginia	10.4	1.65**	1.88**
Washington	18.7	3.66**	3.55**
West Virginia	14.8	1.09**	1.27**
Wisconsin	11.9	0.44	0.34
Wyoming	14.1	0.64	0.99**

**Significantly different from zero at the .05 level, two-tailed test.

Slope" in the table) imposed the restriction that each state share the same slope coefficient for the TUR, whereas the second estimate ("Different Slopes") allowed the coefficients for the TUR to vary among the states.[19] These two estimates are consistent with each other and suggest that, although many states experienced large increases in average UI duration during the mid-1990s, these increases varied significantly in size. For example, 10 states--Connecticut, Delaware, Florida, Hawaii, Maryland, New Hampshire, Oregon, Pennsylvania, Texas, and Washington--experienced increases of about 1.8 weeks or higher under both estimating procedures.[20] The monthly data were examined to understand the reasons for these large state differences.

E. STATE ANALYSIS OF MONTHLY DATA

Monthly data on UI durations at the state level offer both advantages and disadvantages for analysis. Advantages include the large increase in observations because of the use of the monthly data, the ability to examine seasonal patterns at the state level, and the possibility of looking at important subgroups of states.[21] The primary disadvantage of the monthly data is that the number of variables readily available on such a basis is limited. Most important, no monthly data on unemployment durations exist at the state level, so many of the analytic results reported at the national level cannot be duplicated here. Instead, monthly data were used primarily to identify states that appear to have had

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[19]Mirroring the constrained estimates, most state slopes fell in the 0.5-1.0 range.

[20]These states were also estimated to have such increases even after controlling for levels of the four UI-related policy variables described in the text.

[21]The basic data set contains monthly data on 51 UI jurisdictions for the 216-month period 1980 to 1997--a total of more than 11,000 observations.

especially noteworthy trends in average UI duration during the 1993-1997 period.[22] The expansion of observations provided by the use of monthly data also provided the opportunity to measure some of the effects more precisely than was possible with the more aggregated data.

Table II.7 reports descriptive regressions on the monthly data. Other than the TUR and various dummy variables (including the "fixed effects" used in all of the equations), the only other variables for which results are reported are those measuring the percentages of employment in each state that are in the construction and manufacturing industries.[23] Findings for these variables provided some additional insights not available in the more aggregated estimates.

Overall, many of the equations in Table II.7 closely mirror those reported elsewhere. Whenever only the TUR was controlled for, average UI durations were estimated to be about one week higher in the 1993-1997 period than might have been expected (Equation 1). That result tended to persist when the monthly equations were estimated on 12-month moving averages of the underlying data (Equation 4) and in equations that allowed the change in UI duration to vary by state (see Equation 5 and the discussion of Table II.8 that follows).[24,25]

Addition of the industrial composition data to the monthly regressions had a major effect on the estimates of the coefficient for the 1993-1997 dummy variable. The inclusion of these variables reduced the estimated increase in average UI duration in 1993-1997 by approximately half (Equation

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[22]For this analysis, the data set was extended to include 1997 because data for that year are available for all the variables desired.

[23]"Fixed-effects" models include dummy variables for each cross-section entity (here, states).

[24]The weighted average of the state-specific effects from Equation 5 in Table II.7 was 1.04 weeks when average first payments during 1993-1997 were used as weights.

[25]The moving average regressions fit the data much better than did the other regressions. (The standard error in Equation 1 is 2.5 times that in Equation 4.) Even though all of the equations in Table II.7 controlled for the average monthly variation in UI durations across the states, state-specific factors tended to make the UI duration measure volatile on a month-by-month basis.

TABLE II.7

MONTHLY RESULTS FOR ALL STATES

	Dependent Variable: Average UI Duration			12-Month Moving Average	
Independent Variable	Equation 1	Equation 2	Equation 3	Equation 4	Equation 5
Total Unemployment Rate	0.629*** (0.015)	0.642*** (0.018)	0.664*** (0.028)	0.699*** (0.006)	0.683*** (0.006)
1993-1997 Dummy Variable	1.009*** (0.057)	0.483*** (0.068)	—	0.998*** (0.022)	—
Percentage Employed in Construction	—	-37.766*** (3.961)	-52.810*** (6.426)	—	—
Percentage Employed in Manufacturing	—	-15.724*** (1.722)	-20.558*** (2.931)	—	—
State Fixed Effects	Yes	Yes	Yes	Yes	Yes
State-Specific Dummy Variables for 1993-1997	No	No	Yes	No	Yes
R-squared	0.640	0.669	0.677	0.874	0.896
Standard Error	2.477	2.384	2.360	0.950	0.870
F	1,589***	1,292***	292***	5,177***	1,224***

NOTE: All equations also contained 11 monthly dummy variables. For Equation 1, data are from January 1980 to December 1997, and the mean average UI duration is 14.9 weeks. For Equations 2 and 3, data are from January 1983 to December 1997, and the mean average UI duration is 14.9 weeks. For Equations 4 and 5, data are from December 1980 to December 1997, and the mean is 14.0 weeks. Standard errors are in parentheses below the coefficient estimates.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

TABLE II.8

ESTIMATED INCREASE IN AVERAGE UI DURATION FOR 1993-1997

State	Average UI Duration 1980-1992 (Weeks)[a]	Estimated Change for 1993-1997[b]	Percentage Change[c]
Alabama	12.7	0.9	7.3
Alaska	16.1	0.7	4.4
Arizona	14.4	0.6	4.2
Arkansas	13.8	1.6	11.4
California	16.4	0.5	3.0
Colorado	12.9	1.0	7.5
Connecticut	14.8	3.0	20.0
Delaware	15.7	2.2	13.8
District of Columbia	20.5	-0.3	-1.7
Florida	13.4	1.7	12.5
Georgia	10.9	0.2	2.2
Hawaii	14.1	2.2	15.5
Idaho	13.6	0.6	4.2
Illinois	18.6	1.2	6.2
Indiana	13.1	1.0	7.9
Iowa	14.4	0.4	3.0
Kansas	14.7	0.2	1.4
Kentucky	16.2	-0.9	-5.4
Louisiana	17.1	0.1	0.3
Maine	15.3	0.4	2.6
Maryland	16.0	1.2	7.6
Massachusetts	17.4	0.7	4.0
Michigan	16.1	0.3	1.8
Minnesota	16.6	0.6	3.7
Mississippi	13.9	2.1	15.1
Missouri	13.8	1.6	11.7
Montana	14.4	1.5	10.3
Nebraska	12.9	0.2	1.6

TABLE II.8 *(continued)*

State	Average UI Duration 1980-1992 (Weeks)[a]	Estimated Change for 1993-1997[b]	Percentage Change[c]
Nevada	14.3	0.7	5.0
New Hampshire	10.9	2.3	21.0
New Jersey	16.8	1.1	6.4
New Mexico	16.7	0.6	3.6
New York	19.2	0.9	4.8
North Carolina	10.4	1.2	11.5
North Dakota	15.2	-0.6	-4.2
Ohio	16.0	1.1	7.2
Oklahoma	13.8	1.4	10.4
Oregon	15.4	2.1	13.6
Pennsylvania	17.5	1.4	7.8
Rhode Island	16.0	1.0	6.4
South Carolina	11.7	1.0	8.9
South Dakota	12.5	-0.1	-0.8
Tennessee	13.8	0.9	6.2
Texas	14.5	1.5	10.4
Utah	13.7	0.2	1.8
Vermont	15.2	1.4	9.5
Virginia	11.6	1.9	16.6
Washington	16.7	3.0	18.2
West Virginia	16.8	1.0	6.0
Wisconsin	15.0	0.5	3.2
Wyoming	15.2	-0.0	-0.3

[a]Average of monthly measures.

[b]Taken from Table II.7, Equation 5.

[c]Equals (Column 3/Column 2) × 100.

2). An analysis of the trends in the industrial composition of employment suggests that the reduction in the coefficient for the 1993-1997 dummy variable occurred almost completely from inclusion of the manufacturing variable. Overall there was little change in the fraction of employment in construction in the 1993-1997 period compared to earlier periods. Hence, despite the relatively large coefficients reported in Table II.7, changes in construction employment cannot explain recent changes in average UI durations. However, states experienced a reduction of about 2.7 percentage points in the fraction of employment in manufacturing during the 1993-1997 period, compared to historical averages. According to Equation 2, that trend increased average UI duration by between 0.4 and 0.5 week, thereby explaining the decline in the coefficient of the 1993-1997 dummy variable. The manufacturing employment variable in Equations 2 and 3 is, in part, controlling for changing unemployment durations. Because manufacturing unemployment tends to be heavily concentrated in short-term layoffs, a decline in this employment--other things being equal--lengthens UI durations.[26] To the extent that declining manufacturing employment is also associated with increases in worker displacement, the trend can have an additional impact on average UI durations.

To gain additional insights on the relationship between changes in manufacturing employment and UI duration, Table II.8 repeats the state-specific focus of Table II.6, this time using monthly data. In order to mitigate the extreme volatility present in the monthly data, the estimates in Table II.8 are taken from the final moving average regression in Table II.7, Equation 5. The findings closely mirror those in Table II.6, although they focus on the percentage changes during 1993-1997 as departures from the historical (1980-1992) average. Overall six states exhibited increases in average UI duration of more

[26]It should be pointed out, however, that this pattern was not apparent in the national estimates, which found the layoff rate itself to have little influence on UI durations. Hence, the connection between manufacturing employment and UI duration may be more complex than this simple example suggests.

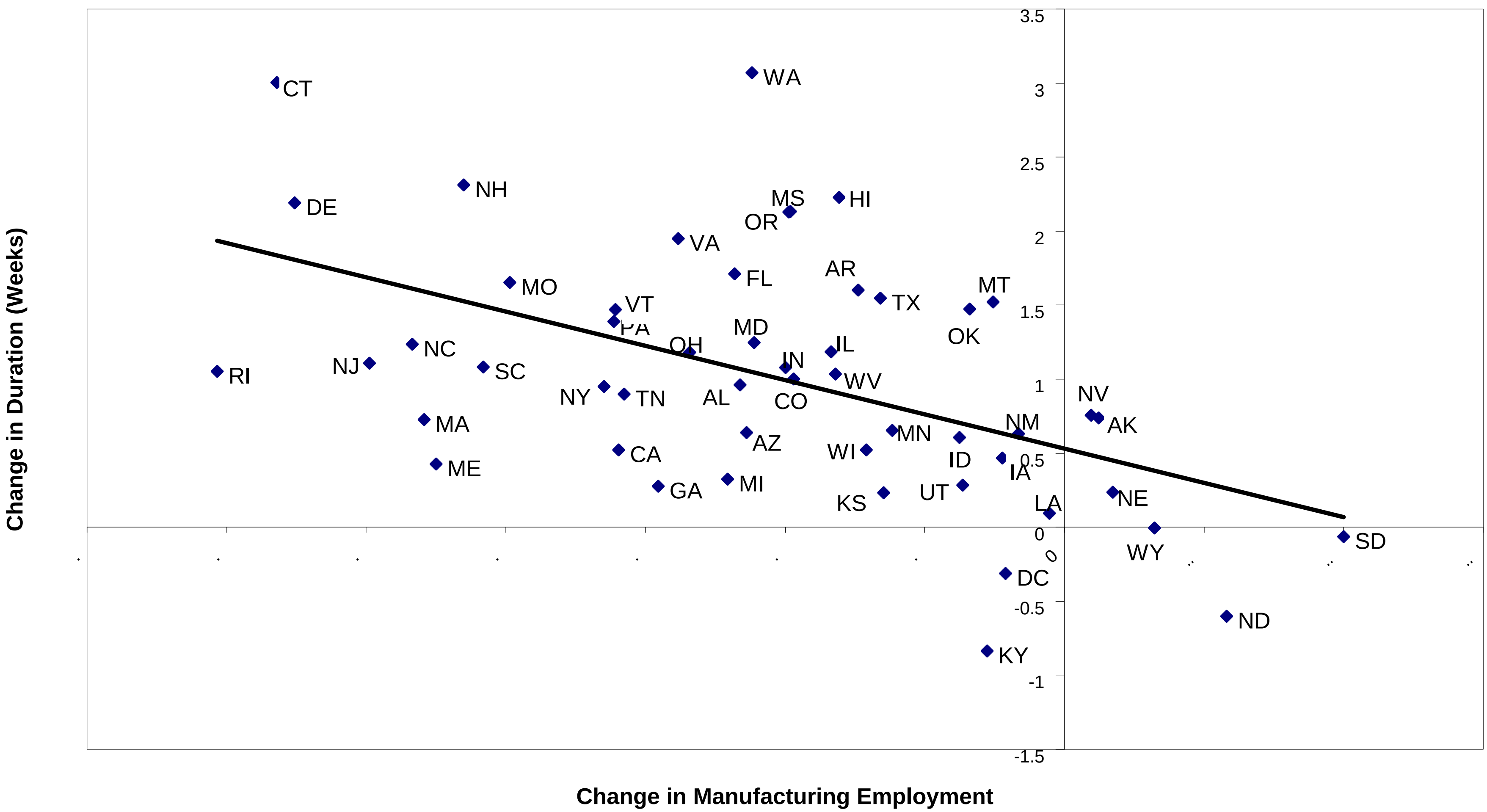
than 15 percent: Connecticut, Hawaii, Mississippi, New Hampshire, Virginia, and Washington. On the other hand, it was estimated that five states exhibited a decline in average UI duration: District of Columbia, Kentucky, North Dakota, South Dakota, and Wyoming. Among the 10 largest states, the variation in average UI durations was somewhat smaller--the largest percentage increase was in Florida (1.7 weeks or about 12.5 percent of the state's pre-1992 average duration), whereas the smallest was in Michigan (0.3 week for a 1.8 percent increase). Some preliminary investigations of these state-specific changes suggest that they may in part be explained by changes in the pattern of employment in the state--especially by the declining importance of manufacturing. An indication of this possibility is provided by the state-specific estimates from the regression that included the construction and manufacturing variables (Equation 3 in Table II.7). They show a quite different pattern than do the estimates in Table II.8. Most notably, the very large estimated increases for Connecticut and Virginia are reduced to statistical insignificance by the inclusion of the industrial composition variables, but those for the other large-increase states are little affected. Similarly, although the largest estimated increase among the top 10 states was in Florida in the unadjusted data, the increase in Florida becomes insignificant once industrial composition is controlled for.[27]

The relationship between changes in manufacturing employment and changes in UI duration is illustrated in Figure II.2. For this figure, the estimated changes in UI duration in each state during the 1993-1997 period (Table II.8) are shown on the vertical axis, whereas estimated changes in the proportion of employment in manufacturing during that period are shown on the horizontal axis.[28] Visually there appears to be some negative slope to this relationship. A simple regression fit to these

[27]Still, the most appropriate state-specific increases in UI duration for some policy purposes are those given in Table III.8, since these estimates reflect benefits that states actually pay (after adjusting for the influence of the business cycle).

[28]The industrial composition variable has been adjusted so as to control for month and the level of the TUR.

Figure II.2
UI Duration and Manufacturing Employment, 1993-1997



data yielded a slope coefficient of -23 (relatively close to the coefficient on manufacturing employment in Equation 3 of Table II.7) with a t-ratio of over 4. Clearly there is some relationship between these two variables. The data for Connecticut, for example, are apparent in the top left quadrant of the figure. The overall clustering of the points in the figure is not tight, however, and some states with major declines in manufacturing employment have not experienced especially large increases in UI duration.[29] Efforts to probe these data further through a series of state-specific regression equations based on the monthly data were largely unsuccessful as many of the coefficients that seemed quite stable in the pooled equations proved to be highly erratic with this disaggregation.

F. CONCLUSIONS

The analysis of the aggregate data therefore allows several conclusions to be drawn about changes in average durations for unemployment insurance benefits during the 1990s.

- ***Average durations of regular UI benefits were between 1.1 and 1.4 weeks longer during the 1993-1996 period than might have been expected on the basis of historical data.*** This finding takes into account the relatively strong labor market prevailing during these years (primarily by controlling for the level of the TUR).

- ***The increase in average durations does not seem to be explained by changes in UI policy at either the state or the federal level.*** It was examined whether increases in average weekly benefits, average potential duration, UI recipiency rates, or the availability of extended benefits could explain the observed increase in UI durations. Although such features of state UI programs as the level of their benefit or the potential duration for which individuals can collect have been found to affect UI durations, these features cannot explain the 1990s increase in average UI duration because these factors did not, in fact, change during the period. Similarly, although availability of extended UI benefits has been found in other studies to increase regular UI durations, such benefits were not available during the mid-1990s in large enough quantities to explain the increase in regular UI durations.

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[29]For example, both New Jersey and Rhode Island had declines in the proportion of employment in manufacturing of about the same size as did Connecticut, but increases in UI duration in these states were slightly below average (about 1.0 week).

- C ***Although attempts to use demographic and economic variables to explain the increase in UI durations at the national level were generally unsuccessful (in part because many of the time series move together), some evidence that suggested that the increase may be arising from changes in the labor market itself was found.*** Specifically, the duration of unemployment itself had a strong effect in explaining UI durations. Once the increase in average unemployment duration during the mid-1990s was controlled for, the estimate for the unexplained increase in UI durations was reduced significantly.

- C ***Analysis of the state-level data suggested that there is an important connection between lengthening UI durations and declining manufacturing employment.*** The ranking of the states by their estimated increases in UI duration was significantly affected when changes in the industrial composition of employment were controlled for, Providing a complete explanation of the connection between these two trends, however, was beyond the limits of the available aggregate data.

III. ANALYSIS OF DATA ON CLAIMANTS

Aggregate measures of UI duration are based on the number of first payments and the total number of weeks of UI collected in a time period. As discussed in Chapter II, these aggregate measures may reflect seasonal patterns in the numerator or denominator but may not adequately relate to the individual experiences of UI claimants. Thus, examination of the experiences of individual claimants may provide an important complement to an analysis of the aggregate trends. Together, these two analytic approaches shed light on the reasons for the change in average UI durations in the 1990s.

The analysis of the claimant-level data is divided into five sections. Section A describes the data used for this analysis, highlighting some of its potential shortcomings for this type of analysis. Section B presents the basic results for the statistical analysis of UI durations using the individual-level data in a state-specific analysis. In Section C, these results are used to assess the quantitative significance of various changes in the composition of UI caseloads in the states. Section D examines equations that pool the data across the states in an effort to better understand how labor market characteristics may have affected the outcomes. Section E determines which industrial groups experience the longest durations, and Section F reviews research that examines how the long-term unemployed support themselves. Finally, Section G summarizes the results.

A. DESCRIPTION OF THE DATA

UI program administrative and survey data on nationally representative random samples of claimants over time would be ideal for analyzing changes in UI durations. With these data, it would be possible to examine the determinants of UI claimants' experiences and how changes in economic circumstances affect such experiences. Unfortunately, such a detailed microdata set is not available, so alternatives were considered.

This project uses more limited data collected as part of two evaluations sponsored by the U.S. Department of Labor (DOL) and conducted by Mathematica Policy Research, Inc.[1] The first data set, collected as part of the Unemployment Insurance Exhaustee Study, contains random samples of claimants in 20 states who collected a UI first payment during a one-year period, from October 1987 to September 1988. These data are fairly representative of UI experiences during periods of relatively healthy labor markets in the 1980s. The second data set, collected as part of the Evaluation of Unemployment Insurance Worker Profiling Initiatives, contains the population of claimants from seven states who collected first payments between July 1995 and December 1996 (except for claimants who were not profiled under the worker profiling and reemployment services system). Again, these data are fairly representative of UI experiences during periods of relatively healthy labor demand--in this case, during the mid-1990s.

It is possible to make only limited comparisons across time periods using these data, for several reasons. First, no survey data, for several reasons, were available from the profiling study. (Administrative and survey data were available from the exhaustee study.) The lack of profiling survey

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[1]Data from another evaluation, the Evaluation of the Emergency Unemployment Compensation Act, are also available. This data set contains random samples of claimants from 16 states who collected first payments between January 1991 and February 1994. However, these data were not included in the analysis, for several reasons. First, the primary analysis question is why the average UI duration did not return to a level that is typical during healthy economic times. A data set on claimants during the recession is less useful in addressing this question compared to the two data sets on claimants during nonrecessionary times. Second, some claimants did not collect regular UI during the recession but collected EUC instead. Since these claimants are not a random subset of all claimants during that period, estimates of average UI duration will be biased. Third, including the EUC data in the analysis would restrict the number of states available for the analysis even further than what is reported here.

data severely limits the ability to assess how claimants' pre-UI job characteristics and reasons for job separation have changed over time.

Second, the sets of states that participated in the evaluations differed. Most important, the seven states in the profiling study were not selected to be nationally representative, so national estimates of UI durations and claimant characteristics cannot be made. Instead, the states used to compare patterns over time are limited to those states that participated in both evaluations.[2] This restriction limits the ability to assess national trends and to make conclusions that apply to the nation in general. Still, it is believed that an examination of the state-specific experiences that is possible with these data can offer insights about the national trends.

Third, the profiling data exclude claimants who were not profiled. The decisions that states used to exclude these claimants could not be perfectly modeled. In general, claimants who were not profiled were likely to be claimants who had definite recall dates to their former employers or who expected to be hired through a union hiring hall. To help ensure comparability across the data sets, claimants in the exhaustees data set who reported they had a definite recall date or who reported they did not search for work because they expected their union to find them a job were excluded.[3] These exclusionary restrictions do not capture all reasons claimants are not profiled, however.[4] Differences in states'

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[2]Data on New Jersey claimants in the mid-1990s were not available in time to be included in the analysis.

[3]The percentage of claimants in the exhaustees data set who were excluded ranged from about 3 percent to about 25 percent. These percentages are low compared to the percentages excluded in the mid-1990s, using comparisons between the number of first payments reported by the states to DOL in the mid-1990s and the sample sizes in the profiling evaluation. This is a particular concern for the Illinois data.

[4]Other common reasons for being screened out before being profiled are that claimants were interstate or transitional claimants, that they had earnings in the first week of benefit collection, and that there was a large gap between when the claimant established a benefit year and when he or she filed for a first week of benefits.

screening procedures for the profiling system may affect the accuracy of these exclusionary restrictions and thus the comparability of the data sets.

Ultimately, four states--Connecticut, Illinois, Mississippi, and Texas--that had relatively comparable information from the two data sources were chosen for the analysis. For most of the analysis, each of these states is examined separately; results for pooling data across the states is mentioned only briefly.

Tables III.1 through III.4 report basic descriptive statistics for the state samples. The tables also indicate whether mean characteristics differ between the 1980s sample (from the study of exhaustees) and the 1990s sample (from the profiling study).[5] However, sample sizes for the exhaustee sample were very small (ranging between 71 and 195, compared both to the number of first payments made in the four states during the late 1980s and to the profiling samples). This suggests that caution should be used in interpreting differences between the two time periods. Still, several trends in the data are consistent across the states. All states experienced relatively large increases in the fraction of UI claimants who are female and the percentage who are African American. Only in Connecticut was there a significant change in the mean age of claimants: in that state, the samples suggest that mean age increased by more than three years. Both Connecticut and Mississippi experienced large declines in the fraction of UI recipients who had been employed in manufacturing.[6] Texas, on the other hand, showed almost no change. In all states, more highly educated individuals were a greater fraction of claimants in the 1990s than in the 1980s. Relatively minor changes were recorded for most of the administrative, UI-related data in the sample

[5]For the exhaustee sample, these means are weighted to adjust for oversampling of UI exhaustees in that study.

[6]Industrial attachment data were unavailable for Illinois in the 1990s sample.

TABLE III.1

DEMOGRAPHIC CHARACTERISTICS, CONNECTICUT
(Standard Deviations in Parentheses)

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Percent Female	36.0 (48.0)	46.6 (49.9)	10.6
Age at First Claim Date (Years)	38.4 (12.5)	41.5 (11.6)	3.1***
Total Unemployment Rate	3.0 (0.1)	5.7 (0.1)	2.7***
Percentage in Manufacturing Industry	35.4 (47.8)	20.5 (40.4)	-14.9***
Race/Ethnicity[a]			
African American	3.8 (19.1)	12.0 (32.5)	8.2***
Asian	3.8 (19.1)	0.1 (0.9)	-3.7***
Caucasian	81.6 (38.8)	77.6 (41.7)	-4.0***
Hispanic	8.3 (27.6)	7.0 (25.4)	-1.3***
Other	0.0 (0.0)	2.7 (16.2)	2.7***
Less than a High School Graduate	28.2	15.3	-12.9***
High School Diploma or GED	42.3	42.0	-0.3***
Vocational/Technical/Some College	17.8	22.8	5.0***
College Degree	8.8	12.5	3.7***
Higher than a College Degree	2.9	7.4	4.5***

TABLE III.1 *(continued)*

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Maximum Benefit Amount (Dollars)	4,769 (144.7)	5,830 (11.0)	1,061***
Base Period Earnings (Dollars)	16,628 (1,372.9)	23,061 (81.7)	6,433***
Weekly Benefit Amount (Dollars)	183.4 (41.4)	229.0 (10.4)	45.6***
Real Maximum Benefit Amount (1980-1982 Dollars)	4,077.5 (123.7)	3,749.3 (7.1)	-328.2***
Real Base Period Earnings (1980-1982 Dollars)	14,224.2 (1,175.8)	14,827.9 (52.5)	603.7
Real Weekly Benefit Amount (1980-1982 Dollars)	156.8 (4.76)	147.2 (0.3)	-9.6**
Potential Duration	26.0 (0.0)	26.0 (0.0)	0
Unweighted Sample Size	**71**	**57,981**	**---**

SOURCE: Data are from the Exhaustees and Profiling evaluations.

[a]Significance levels indicate that the distribution of race/ethnicity differs across the two time periods.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

TABLE III.2

DEMOGRAPHIC CHARACTERISTICS, ILLINOIS
(Standard Deviations in Parentheses)

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Percent Female	38.5 (48.7)	53.3 (49.9)	14.8
Age at First Claim Date (Years)	39.2 (12.8)	37.0 (11.1)	-2.2
Total Unemployment Rate	7.0 (0.1)	5.3 (0.1)	-1.7***
Percentage in Manufacturing Industry	35.3 (47.8)	---	---
Race/Ethnicity[a]			
African American	24.4 (43.0)	28.8 (45.3)	4.4***
Asian	0.4 (6.2)	0.0 (0.0)	-0.4***
Caucasian	63.7 (48.1)	65.1 (47.7)	1.4***
Hispanic	10.7 (31.0)	6.0 (23.7)	-4.7***
Other	0.4 (6.2)	0.1 (3.7)	-0.3***
Less than a High School Graduate	16.1	8.4	-7.7***
High School Diploma or GED	62.9	48.7	-14.2***
Vocational/Technical/ Some College	12.3	23.4	9.1***
College Degree	5.6	14.9	9.3***
Higher than a College Degree	3.1	4.6	1.5***

TABLE III.2 *(continued)*

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Maximum Benefit Amount (Dollars)	3,635 (110.3)	4,818 (7.4)	1,183***
Base Period Earnings (Dollars)	16,706 (1,153.5)	23,048 (111.5)	6,342***
Weekly Benefit Amount (Dollars)	139.8 (41.3)	185.3 (6.6)	45.5***
Real Maximum Benefit Amount (1980-1982 Dollars)	3,117.6 (95.3)	3,094.8 (4.7)	-22.8
Real Base Period Earnings (1980-1982 Dollars)	14,351.6 (999.0)	14,805.5 (71.6)	453.9
Real Weekly Benefit Amount (1980-1982 Dollars)	119.9 (3.7)	119.0 (0.2)	-0.9
Potential Duration	26.0 (0.0)	26.0 (0.0)	0
Unweighted Sample Size	**139**	**54,722**	**---**

SOURCE: Data are from the Exhaustees and Profiling evaluations.

[a]Significance levels indicate that the distribution of race/ethnicity differs across the two time periods.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

TABLE III.3

DEMOGRAPHIC CHARACTERISTICS, MISSISSIPPI
(Standard Deviations in Parentheses)

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Percent Female	44.0 (49.6)	48.9 (50.0)	4.9
Age at First Claim Date (Years)	36.6 (11.6)	36.7 (10.9)	0.1
Total Unemployment Rate	8.6 (0.4)	6.2 (0.2)	-2.4***
Percentage in Manufacturing Industry	63.7 (48.1)	48.9 (50.0)	-14.8**
Race/Ethnicity[a]			
African American	44.6 (49.7)	49.8 (50.0)	5.2***
Asian	0.0 (0.0)	0.2 (4.8)	0.2***
Caucasian	52.8 (49.9)	49.5 (50.0)	-3.3***
Hispanic	2.6 (16.0)	0.3 (5.1)	-2.3***
Other	0.0 (0.0)	0.1 (3.7)	0.1***
Less than a High School Graduate	30.0	21.6	-8.4***
High School Diploma or GED	52.6	53.1	0.5***
Vocational/Technical/Some College	14.1	20.3	6.2***
College Degree	3.3	4.3	1.0***
Higher than a College Degree	0.0	0.8	0.8***

TABLE III.3 *(continued)*

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Maximum Benefit Amount (Dollars)	2,639.3 (114.1)	3,476.1 (5.2)	836.8***
Base Period Earnings (Dollars)	8,449.3 (1,049.1)	14,461.9 (41.8)	6,012.6***
Weekly Benefit Amount (Dollars)	107.6 (27.9)	142.9 (40.9)	35.3***
Real Maximum Benefit Amount (1980-1982 Dollars)	2,264.8 (97.6)	2,236.5 (3.4)	-28.3
Real Base Period Earnings (1980-1982 Dollars)	7,246.3 (902.9)	9,303.3 (26.9)	2,057**
Real Weekly Benefit Amount (1980-1982 Dollars)	92.3 (3.2)	92.0 (0.1)	-0.3
Potential Duration	24.1 (3.6)	23.9 (3.6)	-0.2
Unweighted Sample Size	**97**	**53,299**	**---**

SOURCE: Data are from the Exhaustees and Profiling evaluations.

[a]Significance levels indicate that the distribution of race/ethnicity differs across the two time periods.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

TABLE III.4

DEMOGRAPHIC CHARACTERISTICS, TEXAS
(Standard Deviations in Parentheses)

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Percent Female	32.7 (46.9)	41.9 (49.3)	9.2**
Age at First Claim Date (Years)	38.5 (12.2)	37.8 (11.0)	-0.7
Total Unemployment Rate	7.5 (0.3)	5.7 (0.2)	-1.8***
Percentage in Manufacturing Industry	22.2 (41.6)	22.2 (41.6)	0.0
Race/Ethnicity[a]			
African American	12.9 (33.6)	17.0 (37.5)	4.1***
Asian	0.3 (5.4)	1.1 (10.4)	0.8***
Caucasian	53.7 (49.9)	46.8 (49.9)	-6.9***
Hispanic	29.9 (45.8)	33.9 (47.3)	4.0***
Other	1.6 (12.4)	0.1 (10.8)	-1.5***
Less than a High School Graduate	30.7	23.5	-7.2***
High School Diploma or GED	50.4	45.4	-5.0***
Vocational/Technical/Some College	11.5	19.9	8.4***
College Degree	6.9	7.8	0.9***
Higher than a College Degree	0.5	3.4	2.9***

TABLE III.4 *(continued)*

Variable	Late 1980s	Mid-1990s	Change from Late 1980s to Mid-1990s
Maximum Benefit Amount (Dollars)	3,358.5 (151.3)	4,231.0 (3.8)	872.5***
Base Period Earnings (Dollars)	14,653.0 (1,026.1)	20,171.2 (42.5)	5,518.2***
Weekly Benefit Amount (Dollars)	158.6 (56.0)	192.8 (66.2)	34.2***
Real Maximum Benefit Amount (1980-1982 Dollars)	2,868.4 (129.3)	2,708.8 (2.4)	-159.6
Real Base Period Earnings (1980-1982 Dollars)	12,514.8 (877.2)	12,907.0 (27.1)	392.2
Real Weekly Benefit Amount (1980-1982 Dollars)	135.5 (4.4)	123.4 (0.1)	-12.1***
Potential Duration	20.4 (5.4)	21.1 (4.9)	0.7
Unweighted Sample Size	**195**	**270,666**	**---**

SOURCE: Data are from the Exhaustees and Profiling evaluations.

[a]Significance levels indicate that the distribution of race/ethnicity differs across the two time periods.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

once the data were adjusted for inflation. However, both Connecticut and Texas experienced declines in the average real UI WBA.

In constructing these data sets, TURs were imputed based on the rate that applied for the month in which benefits collection started. According to these data, claimants in three of the states (Illinois, Mississippi, and Texas) experienced substantially stronger labor markets in the 1990s than in the 1980s. In these states, the average unemployment rate declined between 1.7 and 2.4 percentage points from the late 1980s to the mid-1990s. As shown in the previous chapter, such a change would have been expected to reduce UI durations significantly. In Connecticut, on the other hand, measured unemployment rates rose by more than three percentage points--a change that would be expected to increase UI durations substantially.[7]

B. REGRESSION ANALYSIS OF UI DURATION

As discussed in Chapter I, theoretical models suggest that individual characteristics will be important factors that explain differences in the length of unemployment and UI spells. Changes over time in average UI duration can be explained in two ways: (1) by changes in the average characteristics of claimants who collect benefits, and (2) by changes in the effect that these characteristics have on average UI durations. In this section, regression analysis is used to estimate the relative importance of these two factors in explaining the increase in UI durations in the 1990s. Because most of the regression equations were imprecisely estimated for the exhaustee samples, the analysis focuses primarily on assessing the effects of changes in sample characteristics. Changes in the parameters of the model are briefly discussed when pooled results are presented later in this section.

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[7] The aggregate regressions in Chapter II suggest that each percentage point increase in the TUR may increase average UI duration by about 0.7 week. Therefore, other things being equal, average durations in Illinois, Mississippi, and Texas would have been expected to decrease by about 1.4 weeks, whereas average durations in Connecticut might have been expected to increase by about 2.1 weeks.

Table III.5 reports state-specific regressions from the profiling samples.[8] Some preliminary analysis suggested that the appropriate dependent variable in these models was the (natural) logarithm of weeks of UI benefits collected, so all the regressions used that transformation. Thus the individual coefficient estimates in the regressions can be interpreted as the proportional change in UI duration brought about by a one-unit change in the independent variable.

Perhaps the most notable feature of the equations reported in Table III.5 is their overall similarity across the states. For example, in all of the states, women are estimated to collect UI for about five percent more weeks than men.[9] Similarly, African Americans are estimated to collect benefits for between 4 and 30 percent longer than whites (the omitted category in the regressions). Age is also estimated to have a significant positive effect on duration, although here the estimates vary somewhat among the states.[10] In the three states that provided information on industrial

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[8]Ordinary least squares (OLS) regressions are presented in Table III.5. Many different model specifications yielded generally similar results. Among the model specifications tried were quarterly dummy variables, using nominal dollars for the WBA and base period earnings (instead of real dollars), and modeling the level of weeks of benefits collected instead of the logarithm of weeks collected. In addition, models that corrected for censoring of weeks of benefits collected were estimated. Estimates of these models gave similar results. OLS results are presented because they make it easier to interpret the estimates. Finally, estimating the model using the 1980s data and estimating the change in predicted weeks collected between the 1980s and 1990s was tried. These results are not reported, because the smaller sample sizes for the 1980s data led to less stable estimates for the coefficients.

[9]At a mean duration of about 14 weeks, this amounts to 0.7 week.

[10]The significance of the age-squared term, however, indicates that this positive effect eventually becomes negative for older workers.

TABLE III.5

UNEMPLOYMENT INSURANCE DURATION ANALYSIS
(Standard Errors in Parentheses)

Independent Variable	Connecticut	Illinois	Mississippi	Texas
Intercept	1.751*** (0.056)	1.044*** (0.046)	1.487*** (0.061)	1.816*** (0.029)
Female Dummy Variable	0.059*** (0.009)	0.056*** (0.009)	0.067*** (0.010)	0.054*** (0.005)
Age at First Claim (Years)	0.022*** (0.003)	0.044*** (0.002)	0.022*** (0.003)	0.002* (0.001)
Square of Age at First Claim (Years)	-0.0002*** (0.00003)	-0.0004*** (0.00003)	-0.0001*** (0.00004)	-0.00006*** (0.00002)
African American Dummy Variable	0.167*** (0.014)	0.308*** (0.010)	0.040*** (0.010)	0.132*** (0.008)
Hispanic Dummy Variable	0.091*** (0.018)	0.142*** (0.018)	-0.321*** (0.092)	0.004 (0.006)
Less than a High School Diploma or GED	0.026** (0.013)	0.056*** (0.016)	-0.005 (0.012)	0.012* (0.007)
Some College	-0.041*** (0.011)	-0.066*** (0.011)	0.041*** (0.012)	0.010 (0.007)
College Degree	-0.079*** (0.014)	-0.089*** (0.013)	0.026 (0.024)	-0.009 (0.010)
Higher than a College Degree	-0.127*** (0.018)	-0.040* (0.022)	0.066 (0.052)	-0.013 (0.014)
Potential Duration (Weeks)	---	---	-0.001 (0.002)	0.013*** (0.0006)
Manufacturing Dummy Variable	-0.063*** (0.010)	---	-0.067*** (0.010)	-0.264*** (0.006)
Real Weekly Benefit Amount (1980 Dollars x 10)	0.009*** (0.0010)	0.009*** (0.001)	0.027*** (0.002)	0.006*** (0.0008)
Real Base Period Earnings (1980 Dollars x 1,000)	0.0019*** (0.0005)	0.0003 (0.0003)	-0.003*** (0.001)	0.0003 (0.0002)

TABLE III.5 *(continued)*

Independent Variable	Connecticut	Illinois	Mississippi	Texas
R-squared	0.016	0.033	0.022	0.032
F	69.2	153.3	78.7	369.4
Mean of UI Weeks Collected	16.4	12.6	13.8	14.4
Number of Observations	**49,644**	**50,078**	**45,570**	**144,995**

NOTE: Data are from the Profiling Evaluation. The dependent variable is the log of UI weeks collected. Standard errors are in parentheses below the coefficient estimates. Potential duration does not vary in Connecticut and Illinois since they are uniform duration states. Data on industry are unavailable for Illinois.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

attachment in the profiling data, workers from manufacturing industries experienced significantly shorter UI durations than did workers from other industries.[11]

Parameters of the UI system appear to have some influence on duration; because of the nonexperimental nature of these data, however, such estimates should be treated with caution. It is possible that the correlations should more properly be regarded as reflecting relationships between (unmeasured) workers' characteristics and UI parameters rather than true behavioral effects. Still, the equations suggest that potential duration has a positive effect on actual duration in Texas, with each extra week leading to about 0.2 extra week of benefits being collected--a figure close to the consensus econometric estimate. Mississippi, however, exhibits a negative effect of potential duration--a finding difficult to reconcile with the existing literature.[12] Real WBAs also appear to have a positive effect on duration: each $10 increase in the WBA is estimated to increase UI duration by between 0.6 percent (in Texas) and 2.2 percent (in Mississippi). These estimates, although smaller than most of the econometric estimates, suggest that changes in UI benefits could have had some effect on average UI duration.[13]

Modeling labor market effects with the profiling data proved difficult. The variation in unemployment rates within each state during the periods in which the profiling data were collected was

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[11]Dummies for all the one-digit standard industrial classification (SIC) codes were not included in the regressions in Table III.5, because sample sizes for each one-digit SIC code in the data from the late 1980s are extremely small. In addition, use of one-digit standard occupational classification codes in the regressions was explored. These variables were unavailable in Illinois and Mississippi and suffered from similar limitations in Connecticut and Texas.

[12]Because Connecticut and Illinois offer uniform UI durations to their claimants, the potential duration variable was not included in these regressions.

[13]The consensus econometric estimate suggests that each $10 per week should increase UI duration by about five percent (see Decker 1997).

too small to permit accurate or plausible estimates.[14] Therefore, the TUR was not included in the equations in Table III.5. Instead, the information in these regressions was used to evaluate the impact of changes in claimant characteristics in the states, delaying the analysis of the TUR until the pooled regression analysis is discussed.

C. ESTIMATED EFFECTS OF CHANGES IN CLAIMANT CHARACTERISTICS

In this section, the regressions from Table III.5 were used to estimate how the changes in claimant characteristics reported in Section A may have affected UI duration at the state level. Table III.6 shows the basic results from the calculations. Entries in the table indicate the relative importance of various demographic and economic factors in explaining the total changes in average UI duration that can be attributed to such changes. These are most easily understood by starting at the bottom of Table III.6. For example, the entries there show that the profiling regressions, in combination with the data on economic and demographic characteristics of claimants in the 1980s, can explain an increase in average duration in Connecticut of about 7.3 percent. In weeks, that would amount to approximately one week--about one-third of the total three-week increase estimated in Chapter II. Similarly, the estimates would have predicted a *decline* in average duration in Illinois of about 1.4 weeks (versus an estimated increase of 1.2 weeks in Chapter II), an increase of 0.3 week in Mississippi (2.1 in Chapter II), and 0.1 week in Texas (1.5 in Chapter II). Thus, although the ranking of states in Table III.6 in terms of relative changes in average duration is the same as was derived from the aggregate data in Chapter II, the sizes of the absolute predicted changes are much smaller. A substantial portion of the (TUR-adjusted) changes estimated in Chapter II cannot be

[14]The TUR ranged from 5.6 to 6.0 in Connecticut, from 5.0 to 5.4 in Illinois, from 5.8 to 6.5 in Mississippi, and from 5.4 to 6.1 in Texas. Hence, the narrow range of values for the TUR could not be reliably used in the regressions to predict how UI durations would have differed during the late 1980s, when the TUR is out of the sample range for the profiling data.

TABLE III.6

DECOMPOSITION OF CHANGES IN UI DURATIONS
(Change is in Natural Logarithms)

Independent Variable	Connecticut		Illinois		Mississippi		Texas	
	Change (x 100)	Percent of Change	Change (x 100)	Percent of Change	Change (x 100)	Percent of Change	Change (x 100)	Percent of Change
Intercept	0	0	0	0	0	0	0	0
Female Dummy Variable	0.6	8.6	0.8	-8.2	0.3	15.1	0.5	49.8
Age at First Claim (Years)	6.7	92.2	-9.7	94.6	0.2	10.2	-0.2	-17.2
Square of Age at First Claim (Years)	-0.04	-0.5	0.1	-0.9	0.1	3.5	-0.005	-0.5
African American Dummy Variable	1.4	18.7	1.4	-13.2	0.2	9.6	0.5	54.3
Hispanic Dummy Variable	-0.1	-1.6	-0.7	6.5	0.7	33.9	0.01	1.4
Potential Duration	---	---	---	---	0.02	0.8	0.9	91.4
Manufacturing Dummy Variable	0.9	12.8	---	---	1.0	45.4	0.0	0
Real Weekly Benefit Amount (1980 Dollars x 10)	-0.9	-12.3	-0.1	0.9	-0.1	-3.7	-0.7	-75.1
Real Base Period Earnings (1980 Dollars x 1,000)	0.1	1.6	-0.01	0.1	-0.7	-32.0	0.01	1.2
Less than High School Education	-0.3	-4.7	-0.4	4.3	0.05	2.1	-0.1	-8.8
Some Post-Secondary Education	-0.2	-2.9	-0.7	7.3	0.3	11.6	0.1	8.0
College Graduate	-0.3	-4.0	-0.8	8.2	0.02	1.1	0.001	-0.9
Higher than a College Graduate Education	-0.6	-7.8	-0.1	0.6	0.05	2.4	-0.04	-3.6
Total	**7.2978**	**100**	**-10.2060**	**100**	**2.1763**	**100**	**0.9985**	**100**

TABLE III.6 *(continued)*

NOTE: The numbers in this table are constructed from regressions presented in Table III.5. Entries in the "Change" columns represent changes in the natural logarithm of UI weeks collected that are associated with the demographic and UI program characteristics that are explanatory variables in the regressions.

explained by changes in the characteristics of the UI caseload that can be measured with the profiling data.

Despite this underestimation, the individual entries in Table III.6 do offer some intriguing insights about possible reasons for the observed changes. In Connecticut, for example, the majority of the estimated 7.3 percent increase in average duration is attributable to the older average age of recipients in the 1990s. The decline in manufacturing and increased representation of African Americans in Connecticut is also estimated to have had a relatively large effect on the average increase in duration. These older workers may have been permanently displaced from manufacturing jobs as companies shift from temporary layoffs to permanent downsizing of their workforces.

For Illinois, the predicted *decline* in average duration predicted by Table III.6 is anomalous, since the aggregate data suggest otherwise. However, the figures in the table suggest that most of the changes in the demographic profile of Illinois claimants would have predicted an increase in duration, although all of these effects are dominated by the estimated decline in the average age of claimants. These results might have been significantly affected by the absence of the industry variable for the Illinois sample or, possibly, by other data sample quality problems in the profiling data in that state.

The Mississippi results reported in Table III.6 imply that the decline in manufacturing had an important effect on average duration in that state. Demographic factors (such as the increase in the proportion of claimants who are female or African American and the increase in average age) also have had some impact.[15] These have been counterbalanced, however, by increasing average base period

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[15]Because the states may have changed the way they record race/ethnicity and because of the small fraction of claimants in other minority groups besides African American and Hispanic, dummy variables for these groups are not included in the regressions. Inclusion of the other minority categories does not affect the substantive patterns detected.

earnings in Mississippi (other things being equal, average base period earnings tend to reduce average duration).

Finally, the Texas results suggest that the small estimated change in average duration reported in Table III.6 actually reflects offsetting influences of a variety of changes. Tending to increase average duration were changes in the demographic profile of Texas recipients (as in other states, an increase in the representation of female and African American workers) and a modest increase in average potential duration. These effects were offset by a small decline in the average age of Texas recipients and a decline in the real WBA.

Overall, these results, although not strong or dramatic, reinforce the general finding from Chapter II that changes in the average duration of UI claims are arising from the labor market and are being affected by changes in the nature of the UI caseload. However, the microdata (especially the data from the profiling study) are not sufficiently complete (for example, with respect to details on individuals' layoffs) to permit a detailed picture.

D. POOLED ANALYSIS

Initial examination of the data from the exhaustee sample on a state-specific basis yielded statistical results that were highly unstable. Therefore, these regressions could not be compared to the profiling regressions to determine whether some coefficients had changed substantially between the 1980s and 1990s. In this section, the data across all the states are pooled to make such a comparison. Unfortunately, some of the results from the exhaustee data remain anomalous--possibly because of small sample sizes or inadequacies in the weighting schemes that sought to "reverse" the oversampling of exhaustees. The pooled results do offer a few additional insights, however.

Table III.7 reports four regressions that have been pooled across the states. All regressions now contain the TUR, although some concern remains about whether the cross-section differences in TURs observed in these microdata yield estimates that are consistent with the time series cross- section estimates presented in Chapter II. Because the profiling data in Illinois do not contain information on industrial attachment, separate equations for a pooled sample of the three states with such data are presented.

The pooled regressions from the profiling data closely resemble the state-specific regressions reported in Table III.5.[16] Coefficients for many of the demographic variables are close to those reported earlier. For the four-state regression, the coefficient of the TUR is positive and of approximately the same magnitude as obtained in the aggregate regressions in Chapter II.[17] The sign of this coefficient becomes negative, however, in the three-state sample--perhaps because of insufficient cross-section variation in that sample. However, the coefficient for the manufacturing variable in the three-state regression is highly statistically significant, implying that a 10 percentage point decline in the fraction of employment in manufacturing might increase average UI durations by about 2 percent (about 0.3 week).

The pooled exhaustee regressions are much less satisfactory. This may be attributable in part to the much smaller sample sizes (502 in the four-state sample and 363 in the three-state sample), which yield imprecise estimates of various parameters in the model. Some of the earlier findings from the profiling regressions are confirmed in the exhaustee regressions--for example, females and

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[16]Given the large samples available in the profiling data, an F-statistic rejects the hypothesis that the coefficients are identical across the states, however.

[17]The coefficient of .06 implies that each point increase in the TUR increases average duration by six percent. If average durations are about 14 weeks, that would be a 0.84 week increase--a figure close to the 0.7 figure reported for many of the regressions in Chapter II.

TABLE III.7

DURATION ANALYSIS USING POOLED REGRESSIONS
(Standard Errors in Parentheses)

Independent Variable	1990s, Excluding Manufacturing	1990s, Including Manufacturing	1980s, Excluding Manufacturing	1980s, Including Manufacturing
Intercept	1.380***	2.652***	0.458	2.088***
	(0.032)	(0.051)	(0.581)	(0.570)
Female Dummy Variable	0.065***	0.058***	0.233**	0.227**
	(0.003)	(0.004)	(0.103)	(0.103)
Age at First Claim (Years)	0.013***	0.011***	0.026	-0.041
	(0.001)	(0.001)	(0.026)	(0.027)
Square of Age at First Claim (Years)	-0.0001*** (0.00001)	-0.00004*** (0.00001)	-0.0003 (0.0003)	0.0006* (0.0003)
African American Dummy Variable	0.123*** (0.004)	0.100*** (0.005)	0.129 (0.123)	0.055 (0.137)
Hispanic Dummy Variable	0.062***	-0.023***	0.201	0.128
	(0.004)	(0.005)	(0.125)	(0.118)
Manufacturing Dummy Variable	---	-0.193*** (0.004)	---	-0.100 (0.104)
Real Weekly Benefit Amount (1980 Dollars x 10)	0.015*** (0.0004)	0.011*** (0.0005)	0.013 (0.014)	-0.016 (0.014)
Real Base Period Earnings (1980 Dollars x 1,000)	0.0003** (0.0001)	0.0007*** (0.0002)	0.007 (0.006)	0.021*** (0.0006)
Total Unemployment Rate	0.060***	-0.135***	0.136***	0.118***
	(0.005)	(0.008)	(0.042)	(0.034)
R-squared	0.016	0.026	0.040	0.093
F	883.2	765.0	2.6	4.0
Number of Observations	**436,668**	**255,362**	**502**	**363**

NOTE: Data are from the Profiling Evaluation. The dependent variable is the log of UI weeks collected. Standard errors are in parentheses below the coefficient estimates. Regression results for columns that exclude manufacturing contain data for Connecticut, Illinois, Mississippi, and Texas. Regression results for columns that include manufacturing contain data for Connecticut, Mississippi, and Texas.

*Significantly different from zero at the .10 level, two-tailed test.
**Significantly different from zero at the .05 level, two-tailed test.
***Significantly different from zero at the .01 level, two-tailed test.

minority workers tend to collect for longer periods. In addition, the coefficient of the manufacturing dummy is reasonably consistent with that reported for the profiling data. However, examination of the stability of the estimates (not shown) suggests that an evaluation of whether the coefficients are “different” in the 1990s than they were in the 1980s would not be fruitful.

E. DIFFERENCES IN WEEKS COLLECTED, BY INDUSTRY

An examination was made of whether any patterns existed across states in the industries that had long average UI durations (Table III.8).[18] Without adjusting for the other characteristics of workers, two industrial groups--transportation and public utilities and finance, insurance, and real estate--had consistently higher average durations than other groups.[19] Workers in the finance, insurance, and real estate industry group typically are less affected by business cycle downturns, but they experienced a higher-than-usual rate of job separation during the recession in the early 1990s (Gardner 1994). Workers in this group are also less likely than workers in the manufacturing industry to expect to be recalled to their former employers. This pattern is therefore consistent with the literature review in Chapter I, which found that workers who do not expect to be recalled are more likely to experience longer UI spells. In contrast, construction workers and durable manufacturing workers had shorter average durations.

Regression analysis (not shown here) suggests that workers in the following groups were more likely to experience long durations once the other available characteristics were controlled for: agriculture; transportation and public utilities; finance, insurance, and real estate; and public

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[18]Industry data from the profiling time period are not available for Illinois, while sample sizes are too small in the exhaustees data to permit analysis by industry. Analysis of differences in occupational patterns could not be conducted because adequate data are not available.

[19]Several other industries experienced UI durations that were longer or shorter than average, but their patterns were not consistent across states.

TABLE III.8

MEAN UNEMPLOYMENT INSURANCE DURATIONS, BY INDUSTRY
(Weeks)

Industry	Connecticut	Mississippi	Texas
Agriculture, Forestry, and Fishing	16.2	14.6	13.5
Mining	15.8	12.6	15.2
Construction	15.1	13.2	13.9
Durable Manufacturing	16.2	13.6	13.8
Nondurable Manufacturing	16.7	13.3	11.7
Transportation and Public Utilities	18.1	15.0	15.1
Wholesale Trade	16.8	14.8	15.5
Retail Trade	15.9	13.4	14.8
Finance, Insurance, and Real Estate	17.6	16.0	16.1
Services	16.2	14.0	14.5
Public Administration	15.7	15.2	15.8
Number of Observations	**57,981**	**53,299**	**270,666**

NOTE: Data are from the Profiling Evaluation.

administration. Whether this pattern is caused by differences in recall rates or other factors could not be assessed. However, it is important to remember that this analysis is restricted to UI claimants who were screened out before entering the profiling system. It is highly likely that workers who are screened out are concentrated in a few industries; these workers probably have very short average UI durations, so the estimates of average durations reported here are likely to be biased upward for those industries with large concentrations of screened-out claimants.

F. SUMMARY

In this chapter, patterns in UI durations in four states in the mid-1990s were analyzed to determine what claimant characteristics are associated with the increase in average durations during this time. To do so, these data were compared to data from the same states during the late 1980s. Limitations of this analysis stemmed from small sample sizes in the 1980s data, the lack of survey data in the 1990s, and the lack of a nationally representative sample in the 1990s. These limitations prevented an important analysis of the characteristics of the preunemployment jobs and the reasons for job separation. Therefore, the findings from this analysis should be interpreted as suggestive of patterns that should be investigated in a more comprehensive research design that includes data collection on reasons for job separation and the nature of the unemployment spell. Nevertheless, several conclusions support and add to the analysis of aggregate data from Chapter II:

- ***Changes in the composition of claimants can explain a portion of the increase in UI duration.*** Although the changes are not large, shifts may have occurred in the composition of claimants toward groups that are more likely to collect for a long time, such as older workers, females, and African Americans.

- ***Changes in the nature of jobs, such as declines in the prevalence of manufacturing, seem to play an important role in increasing average UI durations.*** Although other research suggests that a decrease in the likelihood of recall among the unemployed may be responsible as well, this pattern could not be investigated with either aggregate or claimant-level data.

- ***Workers from certain industries--such as finance, insurance, and real estate; transportation and public utilities; public administration; and agriculture--seem to experience UI durations that are longer than those experienced by other workers.*** This may occur because of differences in recall rates or other characteristics that could not be measured in detail. However, the relative averages of durations by industry varied across states: in some states, workers from a specific industry may experience long durations, but the opposite may be true in another state.

- ***Changes in UI policy are not responsible for the increase in average UI durations.*** Although the data on these patterns are limited, it appears that WBAs and potential duration of benefits collection have either remained the same or changed in ways that suggest average durations should have decreased, rather than increased, all else equal.

IV. FINDINGS AND LESSONS

The average duration of insured unemployment has remained high since the most recent recession, despite lower unemployment rates. There are several possible explanations for this. They include changes in UI laws, changes in the geographic distribution of claimants, and underlying changes in the composition of the unemployed population. In this evaluation, the literature pertaining to UI durations has been reviewed. Analyses of claimant-level data in four states and aggregate national- and state-level data have also been conducted to investigate the magnitude and sources of the increase in average durations during nonrecessionary times. Although assessing the magnitude of the increase in average durations was tried, this research has been exploratory and could not fully examine all the sources of the increase. Nevertheless, several conclusions from this research can be drawn:

- ***UI durations appear to have increased by between 1.1 and 1.4 weeks in the post-1992 period.*** This is a duration approximately nine percent longer than has historically been the case at this stage in the business cycle. UI durations are both cyclically and seasonally sensitive, so analyses of changes in durations should be conducted with care to account for these influences.

- ***Measures of unemployment duration are crucial in explaining UI duration.*** The most likely explanation for the increase in UI and unemployment durations is a change within the labor market itself, such as increases in demographic groups that are more likely to collect for a long time, like older workers, females, and African Americans or increases in the rate of worker displacement.

- ***Changes in the industrial composition of jobs, such as declines in the prevalence of manufacturing, seem to play an important role in increasing average UI durations.*** Manufacturing employment in general is associated with shorter UI spells, probably because of the greater likelihood of recall (although recall probabilities in our analyses could not be controlled for directly).

- ***Changes in UI policy are not responsible for the increase in average UI durations.*** Although the data on these patterns are limited, it can be concluded that UI recipiency rates,

> WBAs, and the potential duration of benefits collection have either remained the same or changed in ways that suggest average durations should have decreased, rather than increased.

In addition, the review of the literature on long-term UI claimants suggests that income from government transfers, retirement benefits, and other household members' earnings does not significantly increase in response to the unemployment spell. UI benefits are an important component in keeping claimants' households above the poverty line, but they are not often large enough to do so without earnings from other household members.

These findings have several implications. First, these patterns do not appear to be induced by UI policy. The literature review suggests that observed changes in the labor markets are occurring both to workers who claim UI and to those who do not. The UI program can attempt to respond to these changes by improving service delivery to facilitate quick reemployment by workers who are permanently separated from their preunemployment jobs. However, it is not likely to be able to override the labor market changes that determine the composition of the unemployed.

Second, the increase in average durations (which was estimated to be about nine percent) may affect states' abilities to balance their UI trust funds. Although unemployment rates are currently quite low by historical standards, the higher durations suggest that states may be less able to increase their trust fund reserves in anticipation of the next recession than might be expected. When the economy experiences its next cyclical downturn, it is expected that the composition of claimants will change to include more claimants on short-term layoff. However, this phenomenon probably will not outweigh the increase in average durations associated with higher unemployment rates and slack demand for workers. Hence, states may experience increased pressure to raise UI taxes to pay for the additional weeks of benefits that claimants are collecting, on average, or they may be more likely to need to borrow funds to maintain trust fund adequacy. Since the increase in durations is not uniform across states, the magnitude of problems states may experience will also vary, unless effective strategies are taken to alleviate the pressure on the trust fund.

Finally, but possibly most important, these findings suggest that further research is needed in two areas. First, additional information is needed on who displaced workers are, which labor market patterns cause increased rates of permanent job separation, and how policy can most effectively respond to these changes. The U.S. Department of Labor (DOL) is sponsoring a study being conducted by MPR to examine these other related topics using nationally representative data on the characteristics of claimants, the preunemployment jobs, and the reasons for job separation.

Second, additional information is needed to assess the effects of increased durations on UI trust funds. These effects could be simulated using models of state trust funds that project increases in tax payments associated with increases in benefits paid. Research using this strategy has found that trust fund adequacy depends heavily on the tax system's responsiveness to changes in benefits paid (Vroman 1998a). This responsiveness varies widely across states, depending on factors such as the taxable wage base, the relationship between the trust fund balance and the tax schedule in effect, and the percentage of experience-rated employers at the minimum and maximum tax rate schedules. Given the differences across states in their trust fund systems, analysis of the effects of increased average UI durations would probably need to be limited to a few representative states. Use of the Benefit Financing Model maintained by DOL in collaboration with more than 15 states may provide a useful start to this type of analysis.

REFERENCES

Addison, John T., and Pedro Portugal. "The Distributional Shape of Unemployment Duration: A Reply." *Review of Economics and Statistics*, vol. 74, no. 4, November 1992, pp. 717-721.

Addison, John T., and Pedro Portugal. "On the Distributional Shape of Unemployment Duration." *Review of Economics and Statistics*, vol. 69, August 1987, pp. 520-526.

Anderson, Patricia M. "Continuing Eligibility." In *Unemployment Insurance in the United States: Analysis of Policy Issues,* edited by Christopher J. O'Leary and Stephen A. Wandner. Kalamazoo, MI: W.E. Upjohn Institute for Employment Research, 1997.

Baker, M. "Unemployment Duration: Compositional Effects and Cyclical Variability." *American Economic Review*, March 1992, pp. 313-321.

Bassi, Laurie J., and Daniel P. McMurrer. "Coverage and Recipiency." In *Unemployment Insurance in the United States: Analysis of Policy Issues*, edited by Christopher J. O'Leary and Stephen A. Wandner. Kalamazoo, MI: W.E. Upjohn Institute for Employment Research, 1997.

Baumol, William J., and Edward N. Wolff. "How Technological Change Increases the Duration of Unemployment." Public Policy Brief no. 41A. Annadale-on-Hudson, NY: The Jerome Levy Economics Institute of Bard College, 1998.

Blanchard, Olivier. "Consumption and the Recession of 1990-1991." *American Economic Review*. Papers and Proceedings of the Hundred and Fifth Annual Meeting of the American Economic Association, Anaheim, CA, January 5-7, 1993, vol. 83, no. 2, May 1993, pp. 270-274.

Blank, Rebecca M., and David E. Card. "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" *Quarterly Journal of Economics*, November 1991, pp. 1157-1189.

Boisjoly, Johanne, and Greg J. Duncan. "Job Losses Among Hispanics in the Recent Recession." *Monthly Labor Review*, vol. 117, no. 6, June 1994, pp. 16-23.

Brewster, J. Alan, Walter Corson, Walter Nicholson, John Friedmann, and Andrea Vayda. "Final Report: Follow-Up Study of Recipients of Federal Supplemental Benefits." Princeton, NJ: Mathematica Policy Research, Inc., September 1978.

Burdett, Kenneth. "Unemployment Insurance Payments as a Search Subsidy: A Theoretical Analysis." *Economic Inquiry*, vol. 17, no. 3, 1979, pp. 333-343.

Burdett, Kenneth, and Jan I. Ondrich. "How Can Changes in Labor Demand Affect Unemployed Workers?" *Journal of Labor Economics*, vol. 3, January 1985, pp. 1-10.

Butler, Richard J., and James B. McDonald. "Trends in Unemployment Duration Data." *Review of Economics and Statistics*, vol. 68, no. 4, November 1986, pp. 545-557.

Corson, Walter, Paul T. Decker, Shari Miller Dunstan, and Anne R. Gordon. "The New Jersey Unemployment Insurance Reemployment Demonstration Project: Final Evaluation Report." Unemployment Insurance Occasional Paper 89-3. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1989.

Corson, Walter, and Mark Dynarski. "A Study of Unemployment Insurance Recipients and Exhaustees: Findings from a National Survey." Unemployment Insurance Occasional Paper 90-3. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1990.

Corson, Walter, Jean Grossman, and Walter Nicholson. "An Evaluation of the Federal Supplemental Compensation Program." Unemployment Insurance Occasional Paper 86-3. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1986.

Corson, Walter, David Horner, Valerie Leach, Charles Metcalf, and Walter Nicholson. "A Study of Federal Supplemental Benefits and Special Unemployment Assistance." Princeton, NJ: Mathematica Policy Research, Inc., 1977.

Corson, Walter, Karen Needels, and Walter Nicholson. "Emergency Unemployment Compensation: The 1990s Experience." Princeton, NJ: Mathematica Policy Research, Inc., January 1999.

Corson, Walter, and Walter Nicholson. "An Examination of Declining UI Claims During the 1980s." Unemployment Insurance Occasional Paper 88-3. Washington, DC: U.S. Department of Labor, September 1988.

Corson, Walter, and Walter Nicholson. *The Federal Supplemental Benefits Program: An Appraisal of Emergency Extended Unemployment Insurance Benefits*. Kalamazoo, MI: The W.E. Upjohn Institute for Employment Research, 1982.

Decker, Paul T. "Work Incentives and Disincentives." In *Unemployment Insurance in the United States: Analysis of Policy Issues*, edited by Christopher J. O'Leary and Stephen A. Wandner. Kalamazoo, MI: W. E. Upjohn Institute for Employment Research, 1997.

Decker, Paul T., and Christopher J. O'Leary. "Evaluating Pooled Evidence from the Reemployment Bonus Experiments." *Journal of Human Resources*, vol. 30, no. 3, 1995, pp. 534-550.

Dynarski, Mark, and Steven M. Sheffrin. "The Behavior of Unemployment Durations Over the Cycle." *Review of Economics and Statistics*, vol. 72, no. 2, 1990, pp. 350-356.

Farber, Henry S. "Are Lifetime Jobs Disappearing? Job Duration in the United States, 1973-93." In *Labor Statistics Measurement Issues*, edited by John Haltiwanger, Marilyn Manser, and Robert Topel. Chicago: University of Chicago Press, 1998.

Farber, Henry S. "The Changing Face of Job Loss in the United States, 1981-95." *Brookings Papers on Economic Activity: Microeconomics*, 1997a.

Farber, Henry S. "Trends in Long-Term Employment in the United States, 1979-96." Industrial Relations Section Working Paper no. 384. Princeton, NJ: Princeton University, July 17, 1997b.

Flinn, C.J., and J.J. Heckman. "Models for the Analysis of Labor Force Dynamics." In *Advances in Econometrics*, edited by R. Basmann and G. Rhodes. Greenwich, CT: JAI Press, 1982, pp. 35-95.

Gardner, Jennifer M. "The 1990-91 Recession: How Bad Was the Labor Market?" *Monthly Labor Review*, vol. 117, no. 6, June 1994, pp. 3-11.

Greene, William H. *Econometric Analysis*. Second edition. New York: Macmillan Publishing Company, 1993.

Grossman, Jean Baldwin. "The Work Disincentive Effect of Extended Unemployment Compensation: Recent Evidence." *Review of Economics and Statistics*, vol. 61, February 1989, pp. 159-164.

Hall, Robert E. "Lost Jobs." *Brookings Papers on Economic Activity*, vol. 1, 1995, pp. 221-256.

Hall, Robert E. "Macro Theory and the Recession of 1990-1991." *American Economic Review*. Papers and Proceedings of the Hundred and Fifth Annual Meeting of the American Economic Association, Anaheim, CA, January 5-7, 1993, vol. 83, no. 2, May 1993, pp. 275-279.

Hall, Robert E. "The Importance of Lifetime Jobs in the U.S. Economy." *American Economic Review*, vol. 72, September 1982, pp. 716-724.

Hansen, Gary D., and Edward C. Prescott. "Did Technology Shocks Cause the 1990-1991 Recession?" *American Economic Review*. Papers and Proceedings of the Hundred and Fifth Annual Meeting of the American Economic Association, Anaheim, CA, January 5-7, 1993, vol. 83, no. 2, May 1993, pp. 280-286.

Ilg, Randy E. "Long-Term Unemployment in Recent Recessions." *Monthly Labor Review*, vol. 117, no. 6, June 1994, pp. 12-15.

Katz, Larry. "Layoffs, Recall, and the Duration of Unemployment." NBER Working Paper no. 1825. Cambridge, MA: National Bureau of Economic Research, January 1986.

Katz, L.F., and B.D. Meyer. "The Impact of the Potential Duration of Unemployment Benefits on the Duration of Unemployment." *Journal of Public Economics*, February 1990, pp. 45-72.

Kletzer, Lori G. "Job Displacement." *Journal of Economic Perspectives*, vol. 12, no. 1, winter 1998, pp. 115-136.

Levenson, Alec R. “Recent Trends in Part-Time Employment.” *Contemporary Economic Policy*, vol. 14, no. 4, October 1996, pp. 78-89.

Levine, Phillip B. “Spillover Effects Between Insured and Uninsured Unemployed.” *Industrial and Labor Relations Review*, vol. 47, no. 1, 1993, pp. 73-86.

Lilien, David. “Sectoral Shifts and Cyclical Unemployment.” *Journal of Political Economy*, vol. 90, 1982, pp. 777-793.

Loungani, P., and B. Trehan. “Explaining Unemployment: Sectoral vs. Aggregate Shocks.” *Federal Reserve Bank of San Francisco Economic Review*, 1997, pp. 3-15.

McMurrer, Daniel P., and Amy B. Chasanov. “Trends in Unemployment Insurance Benefits.” *Monthly Labor Review*, vol. 118, no. 9, September 1995, pp. 30-38.

Moffitt, Robert, and Walter Nicholson. “The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits.” *Review of Economics and Statistics*, vol. 64, no. 1, 1982, pp. 1-11.

Mortensen, Dale T. “Unemployment Insurance and Job Search Decisions.” *Industrial and Labor Relations Review*, vol. 30, no. 4, July 1977, pp. 505-517.

Portugal, Pedro, and John T. Addison. “Problems of Sample Construction in Studies of the Effect of Unemployment Insurance on Unemployment Duration.” *Industrial and Labor Relations Review*, vol. 43, April 1990, pp. 463-477.

Rogers, Cynthia L. “Expectations of Unemployment Insurance and Unemployment Duration.” *Journal of Labor Economics*, vol. 16, no. 3, July 1998, pp. 630-666.

Smith, Ralph E., and Bruce Vavrichek. “Family Incomes of Unemployment Insurance Recipients and the Implications for Extending Benefits.” Washington, DC: Congress of the United States, Congressional Budget Office, February 1990.

Solon, Gary. “Work Incentive Effects of Taxing Unemployment Benefits.” *Econometrica*, vol. 53, no. 2, March 1985, pp. 295-306.

Spiegelman, Robert, and Stephen A. Woodbury. “Bonuses to Workers and Employers to Reduce Unemployment: Randomized Trials in Illinois.” *American Economic Review*, vol. 77, no. 4, 1987, pp. 513-530.

Swaim, P., and M. Podgursky. “The Distributional Shape of Unemployment Duration: A Reconsideration.” *Review of Economics and Statistics*, vol. 74, no. 4, November 1992, pp. 712-717.

U.S. Department of Labor. "A Dialogue: Unemployment Insurance and the Employment Service Programs." Washington, DC: DOL, 1998.

Valletta, Robert G. "Has Job Security in the U.S. Declined?" *Federal Reserve Bank of San Francisco Economic Letter*, no. 96-07. February 16, 1996.

Van den Berg, Gerard J. "The Effects of Changes of the Job Offer Arrival Rate on the Duration of Unemployment." *Journal of Labor Economics*, vol. 12, July 1994, pp. 478-498.

Vroman, Wayne. "Implementing ABP: Impact on State Agencies, Employers, and the Trust Fund." Unemployment Insurance Occasional Paper 98-4. Washington, DC: U.S. Department of Labor, 1998.

Vroman, Wayne. "Labor Market Changes and Unemployment Insurance Benefit Availability." Unemployment Insurance Occasional Paper 98-3. Washington, DC: U.S. Department of Labor, 1998.

Vroman, Wayne. "The Decline in Unemployment Insurance Claims Activity in the 1980s." Unemployment Insurance Occasional Paper 91-2. Washington, DC: U.S. Department of Labor, 1991.

Woodbury, Stephen A., and Murray Rubin. "The Duration of Benefits." In *Unemployment Insurance in the United States: Analysis of Policy Issues*, edited by Christopher J. O'Leary and Stephen A. Wandner. Kalamazoo, MI: W. E. Upjohn Institute for Employment Research, 1997.

APPENDIX A

DATA SERIES AND VARIABLE DEFINITIONS

TABLE A.1

ACTUAL AND PREDICTED AVERAGE UI DURATIONS
1978 to 1996

Year	Average UI Duration	Predicted UI Duration 1	Predicted UI Duration 2
1978	13.3	14.3	13.6
1979	13.1	14.1	13.3
1980	14.9	15	14.1
1981	14.4	15.3	14.7
1982	15.9	16.6	16
1983	17.5	16.5	17.1
1984	14.4	15.2	15.8
1985	14.2	15	15
1986	14.5	14.9	14.8
1987	14.6	14.3	14.3
1988	13.7	13.9	13.8
1989	13.2	13.8	13.3
1990	13.4	14	13.5
1991	15.4	14.8	14.4
1992	16.2	15.2	15.7
1993	15.9	14.8	15.5
1994	15.5	14.3	15.4
1995	14.7	14	14.6
1996	14.9	13.9	14.6

SOURCE: Data on actual duration are from the ETA394. Predicted values are the authors' calculations using data from the ETA394 and the BLS web site.

NOTE: These numbers are shown graphically in Figure II.1. Model 1 uses the total unemployment rate to predict UI duration. Model 2 uses the total unemployment rate and average unemployment duration.

TABLE A.2

DEFINITION OF NATIONAL VARIABLES USED IN ANNUAL ANALYSIS
(1971 to 1996)

Variable	Mean	Standard Deviation
Average Duration of Regular UI Benefits, in Weeks[a]	14.72	1.16
Total Unemployment Rate, Civilians 16 or Older[b]	6.77	1.28
Dummy Variable, = 1 for Years 1993 to 1996	0.21	0.42
Duration of Unemployment, in Weeks[b]	15.06	2.71
Proportion of Total Employment in Construction[b]	0.06	0.003
Proportion of Total Employment in Manufacturing[b]	0.23	0.033
Proportion of Total Employment Who Are Females[b]	0.45	0.024
Average Potential Duration for Regular Unemployment Benefits, in Weeks[a]	24.03	0.23
Ratio of Average UI Benefit to Average Weekly Wage[a]	0.36	0.007
Ratio of Insured Unemployment Rate to Total Unemployment Rate[a]	0.42	0.048

[a]Data from UI Database form ETA394.

[b]Data from DRI National Database.

TABLE A.3

DEFINITION OF NATIONAL VARIABLES USED IN QUARTERLY ANALYSIS
(First Quarter of 1991 to Fourth Quarter of 1996)

Variable	Mean	Standard Deviation
Average Duration of Regular UI Benefits, in Weeks[a]	15.19	2.13
Total Unemployment Rate, Civilians 16 or Older[b]	6.8	1.29
Dummy Variable, =1 for Years 1993 to 1996	0.15	0.36
Growth Rate of Real GDP (annualized)[b]	2.76	3.75
Rate of Capacity Utilization – All Industries[b]	81.5	3.51
Duration of Unemployment, in Weeks[b]	14.37	2.86
Proportion of Unemployment over 27 Weeks[b]	0.14	0.05
Ratio of Insured Unemployment Rate to Total Unemployment Rate[a]	0.49	0.1

[a]Data from UI Database.

[b]Data from DRI National Database.

TABLE A.4

DEFINITION OF NATIONAL VARIABLES USED IN MONTHLY ANALYSIS
(January 1971 to December 1996)

Variable	Mean	Standard Deviation
Average Duration of Regular UI Benefits, in Weeks[a]	15.2	2.77
Total Unemployment Rate, Civilians 16 or Older[b]	6.69	1.29
Dummy Variable, = 1 for Years 1993 to 1996	0.15	0.36
Duration of Unemployment, in Weeks[b]	14.37	2.86
Proportion of Unemployed over 27 weeks[b]	0.14	0.05
Layoffs as a Proportion of Total Labor Force[b]	0.01	0.004
Ratio of Insured Unemployment Rate to Total Unemployment Rate[a]	0.49	0.1

[a]Data from UI Database.

[b]Data from DRI National Database.

TABLE A.5

DEFINITION OF STATE VARIABLES USED IN ANNUAL ANALYSIS
(1978 to 1996)

Variable	Mean	Standard Deviation
Average Duration of Regular UI Benefits, in Weeks[a]	13.82	2.68
Total Unemployment Rate, Civilians 16 or Older[b]	6.51	2.11
Dummy Variable, = 1 for Years 1993 to 1996	0.21	0.41
Average Potential Duration for Regular Unemployment Benefits, in Weeks[a]	23.61	2.36
Ratio of Average UI Benefit to Average Weekly Wage[a]	0.37	0.05
Ratio of Insured Unemployment Rate to Total Unemployment Rate[a]	0.43	0.13
Dummy Variable for Years in Which FSC or EUC Benefits Were Available	0.42	0.49

[a]Data from UI Database.

[b]Data from BLS Homepage.

TABLE A.6

DEFINITION OF STATE VARIABLES USED IN MONTHLY ANALYSIS
(January 1980 to December 1997)

Variable	Mean	Standard Deviation
Average Duration of Regular UI Benefits, in Weeks[a]	14.93	4.1
Total Unemployment Rate, Civilians 16 or Older[b]	6.41	2.2
Proportion of Employment in Construction[b]	0.047	0.012
Proportion of Employment in Manufacturing[b]	0.166	0.069
Dummy Variable, =1 for Years 1993 to 1996	0.31	0.46

[a]Data from UI Database.

[b]Data from BLS Homepage.